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STABILIZATION, PERSISTENCE AND INFLATIONARY CONVERGENCE: A COMPARATIVE ANALYSIS

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Resumen

El objetivo de este artículo es analizar desde un punto de vista empírico la persistencia inflacionaria de un grupo de dieciséis países desarrollados y en desarrollo. Estudiamos en gran detalle las propiedades seriales de la inflación usando métodos univariados y multivariados. Mostramos que la inercia inflacionaria varía bastante tanto entre países como durante el tiempo en un mismo país. Es más, parece haber una clara conexión entre los niveles estimados de inercia y los niveles inflacionarios de los países analizados. Estabilizaciones basadas en un ancla cambiaria parecen reducir la inercia inflacionaria transitoriamente hasta que el tiempo pasa y el programa no logra reducir la inflación a niveles internacionales.

Abstract

The purpose of this paper is to analyze empirically the importance of inflationary persistence in a group of sixteen advanced and developing countries. We do this by studying, in great detail, the time series properties of inflation using both univariate as well as multivariate techniques. We show that inflationary inertia greatly varies both across countries and across different periods of time within a country. Moreover, there seems to exists a clear connection between our estimates of inertia and the levels of inflation in the countries in our sample. Nominal exchange rate anchors seem to reduce inflation inertia but only until time goes by and the program fails to lower inflation to international levels.

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*Fernando Lefort was at the Central Bank of Chile while this paper was written.

I. Introduction

The role of inflationary persistence has become a recurrent theme in discussions on stabilization programs (Fischer 1986). This has been particularly the case in recent debates on the merits of exchange-rate based stabilization plans. Some authors have claimed that anti inflationary programs based on predetermined nominal exchange rates are effective in reducing inflation; others have argued, however, that in the context of stubborn inflationary inertia these type of programs are bound to generate real exchange rate overvaluation and, in some cases, can even result in major crises.[1]

Discussions on the selection of exchange rate regimes in developing countries and transitional economies have also dealt with persistence issues. For example, critics of crawling peg-based exchange rate systems -- including backward-looking exchange rate bands, such as the one operating in Chile since1984 -- have argued that this type of system tends to generate a very high degree of inflationary inertia that can, in the presence of shocks to fundamentals, become highly destabilizing. Those that favor a fixed exchange rate system, on the other hand, have argued that one of the most important merits of this regime is that it will tend to generate a rapid (almost instantaneous) elimination of inflationary inertia, helping stabilize prices at a relatively low cost in terms of output.

In his pioneering work on chronic inflationary experiences in four Latin American countries, Pazos (1969) pointed out that backward looking wage indexation is bound to generate a high degree of inertia, serious destabilizing forces, and a costly and protracted inflationary process.[2]

More recently, there has been increasing interest in analyzing inflationary

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[1] See, for example, Bruno (1993), Dornbusch and Warner (1994) and Edwards (1993). In this paper we use the terms persistence and inertia interchangeably.

persistence in the context of staggered contracts models of the U.S. economy. Fuhrer and Moore (1995), for example, have argued that standard Taylor (1980)-type models fail to generate the degree of inflationary persistence observed in the U.S. data. As an alternative they propose a model where wage setters care about relative real wages over the length of the contract. They show that in this case inflation will show considerable persistence over time. Their empirical analysis --based on autocorrelations and impulse response functions -- supports their model and shows that in the period 1965-1991 inflation in the US has indeed experienced a considerable degree of inertia. As most studies on the subject, however, Fuhrer and Moore (1995) assume that the degree of inflationary inertia in the U.S. has been invariant throughout the period under study. There is no reason, of course, for this to be the case. Indeed, one would expect that as the average length of contracts and the expectations formation process change, so will the degree of inertia. A changing degree of inflationary persistence through time would indeed have important policy implications, especially regarding the output costs of disinflationary programs.[3]

The purpose of this paper is to analyze empirically the importance of inflationary persistence in a group of sixteen advanced and developing countries. We do this by studying, in great detail, the time series properties of inflation using both univariate as well as multivariate techniques. We use monthly data from 1970 through the end of 1996, and we ask the following questions:

- What has been the degree of persistence, if any, in the inflationary processes?
- Has inflation exhibited, in any of the countries in the sample, a unit root?
- Has inertia been stable through time, or has it experienced significant changes?

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[2] See calvo 1983, Fischer 1986 and Taylor 1980 for inertial models based on staggered contracts.
[3] On the costs of disinflation see, for example, Ball (1991).

- In those countries where inertia has exhibited changes, have these been associated to specific policy actions, including the implementation of exchange-rate based stabilization programs?
- Has the degree of inertia been associated to the level of inflation?
- Is there preliminary evidence suggesting that the degree of inflationary inertia has been associated to institutional arrangements governing the existence of indexation in the labor and/or financial markets?

The paper is organized as follows: Section I is the introduction. Section II provides a brief discussion on alternative ways of measuring inertia. In section III we present a preliminary data analysis. We begin the section by providing a brief description of the data set and explain the criteria used for selecting the sample. We then present results from our analysis of inflationary persistence using a full sample (1970-1994). In section IV we provide a detailed comparative analysis of inflationary inertia in 16 countries. We perform a battery of tests: (a) Rolling Augmented Dickey-Fuller tests and (b) Sequential Perron tests to test the unit-root hypothesis in the presence of structural breaks, and (c) Sequential Quandt likelihood ratio tests to analyze the instability of the autoregressive coefficients. We then turn to the estimation of rolling measures of persistence. For all countries we estimate rolling AR(1) coefficients, and report rolling estimates of the dominant root from rolling AR(2) estimates. We then provide an analysis of the connection between our estimates of inertia and the levels of inflation in our countries. In section V we provide a discussion on the evolution of inflationary inertia in three countries that have relied, at one point or another, on a nominal exchange rate anchor to achieve price stability: Chile, Israel and Mexico. Our interest here is to investigate the extent to which these programs have reduced the degree of inflationary inertia. We close in section VI with concluding remarks.

II. Measuring Inflationary Persistence: Analytical Issues

A central purpose of this paper is to provide simple and reliable measures of inflationary persistence for a set of countries.[4] There is by now a large literature, both theoretical and applied, on the persistence displayed by economic time series. Applications include the analysis of U.S. real GDP, commodity prices and inflation.

The measurement of inflationary persistence can be approached either from the perspective of a particular theoretical model, or following a model-free strategy. While the model-based approach has some clear advantages – including its elegance and the existence of sharp hypotheses -- , it also has some limitations. Obstfeld (1995), for example, has persuasively argued that model-based measures of inertia are overly restricted and are subject to mispecification problems that can result in serious biases. In this paper we follow Obstfeld's recommendations and analyze the degree of inflationary inertia from a model free perspective – see, however, Edwards (1996, 1998) for inflationary analyses based on credibility models.

In general, there are two levels in identifying the degree of persistence displayed by a particular time series. On the one hand, we could be interested in determining whether a series is covariance-stationary. The reason for this is that in a stationary series the effect of a shock eventually dies out. In other words the impulse response is zero for a T large enough. In contrast, non-stationary time series exhibit very high persistence, with all shocks qualifying as permanent ones. On the other hand, within the class of stationary time series, some may exhibit higher persistence in the sense that will take longer for the effect of the shock to wear off. Measures like the largest autoregressive root, the half-life and the

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[4] As pointed out above, in this paper we indistinctively use the terms "inflationary inertia" and "inflationary persistence" . See the discussion in this section for formal definitions.

impulse-response function have been used elsewhere for this purpose. In this paper we consider two approaches to measure inertia.

There are several ways of finding out whether a series is non-stationary. Consider the following univariate process.

(1) $y_t = m + a(L)e_t$

(2) $a(L)e_t = e_t + a_1 e_{t-1} + a_2 e_{t-2} + \text{L}$

where $\sum_{j=0}^{\infty} |a_j| < \infty$, the roots of $a(z) = 0$ lie outside the unit circle, and $\{e_t\}$ is a white noise sequence with mean zero and variance s^2. The process described by equations (1) and (2) is a covariance-stationary process. In terms of the persistence displayed by the process, the most interesting property is that the dynamic multiplier (or impulse response function) goes to zero as time goes on,

(3) $\lim_{s\to\infty} \frac{\P y_{t+s}}{\P e_t} = 0.$

That is, the effect of a shock eventually dies off.

Consider, instead, the following unit root process,

(4) $(1-L)y_t = d + b(L)e_t$

where it is required that $b(1) \neq 0$.[5] In this case, an innovation in e_t has a permanent effect on y_t. To see this notice that

$$\frac{\P y_{t+s}}{\P e_t} = 1 + b_1 + b_2 + \text{L} \; + b_s$$

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[5] To understand why is that this condition is required for non-stationarity of y_t, notice that the stationary process in (1) could be first differenced and re-written,

$$(1-L)y_t = (1-L)a(L)e_t = b(L)e_t.$$

Because y_t is stationary, then so is Δy_t. Although this process has the same form as (4), it is clear that necessarily $b(1) = 0$.

and therefore,

$$(5)\ \lim_{s\to\infty}\frac{\partial y_{t+s}}{\partial e_t}=b(1).$$

There have been three main approaches to detect extreme persistence (non-stationarity) in economic time series. Campbell and Mankiw (1987) used a parametric approach to analyze the time series properties of real GNP. Their test makes use of the fact that a non-stationary process requires $b(1)\neq 0$, by calculating impulse-response functions obtained from the estimation of ARMA representations of the first-differenced series. A second approach due to Cochrane (1988) provides a non-parametric way of assessing the same question. Intuitively, Cochrane's test, known as the variance-ratio test, exploits the fact that if y_t is a random-walk ($b(L)=1$) the variance of its s-differences grows linearly with s, while if it is stationary this variance approaches a constant.[6] Consequently, the variance ratio is simply,

$$(6)\ VR(s)=\frac{1}{s+1}\frac{V(y_{t+s+1}-y_t-ms)}{V(y_{t+1}-y_t-m)}.$$

And this ratio approaches 1 for a random-walk and zero for a stationary time series. It is easy to show that in the general case of a non-stationary time series the square root of *VR(s)* would be a lower bound of $b(1)$.

A third approach to measure persistence is known as the unobserved components approach. In this method, the series is modeled as the sum of two components. The best known decomposition is due to Beveridge and Nelson and separates the series in a random

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[6] More generally, let y_t be a non-stationary process of the form indicated in equation (4). Define
$V(s)=E(y_{t+s}-y_t-ms)$,
then

walk and a stationary component.

$$\begin{aligned} (7)\quad & y_t = z_t + c_t \\ & z_t = m + z_{t-1} + h_t \\ & c_t = g(L)d_t \end{aligned}$$

where $E(h_t d_t)$ is arbitrary. This approach measures persistence as the variance of the random walk component relatively to the total variance, $\frac{s^2_{\Delta z}}{s^2_{\Delta y}}$.

These three measures of persistence are simply different ways of scaling the spectral density of Δy_t at zero frequency (Pesaran et al 1993),. This last function can be efficiently estimated using the Barlett kernel,

$$(8)\quad \hat{s}_{\Delta y}(0) = \frac{1}{2p}\left\{\hat{g}_0 + 2\sum_{j=1}^{q}\left[1 - \frac{j}{(q+1)}\right]\hat{g}_j\right\}$$

These three approaches are subject to an important limitation, however,[7] since for any non-stationary general process there exists some stationary process, impossible to distinguish from the former for any finite simple of size T. A way of circumventing this problem and increasing the power in testing for unit root processes is to restrict oneself to assume that the series is well represented by a low order autoregressive process. Although, in general, it is not possible to refute the hypothesis that a general process is a unit-root process, it is generally possible to test the less general hypothesis that a process is an AR(p) process containing a unit-root. The most popular unit root tests among applied economists such as the Dickey-Fuller and Phillips-Perron belong to this family, and assume that the series is represented by a first order autoregressive process.

$$\lim_{s\to\infty}\frac{V(s)}{s} = s^2[b(1)]^2 .$$

[7] See Hamilton (1994) for a discussion in the observational equivalence problem.

A different way of thinking about inertia is within the class of stationary processes. Notice, that the stationarity of an ARMA(p,q) model depends on the roots of the autoregressive part lying outside the unit circle, regardless of the roots of the moving average part. This result provides a simple way of measuring the degree of persistence of a stationary series. In the long run, the dynamics of a series will be dominated by the largest root of its autoregressive part. Of course, if the series is stationary, with the modulus of all inverted roots being less than one, the effect of a random disturbance will eventually die off. In the meantime, however, the effect of the shock will last longer the larger is the value of the inverted dominant root or eigenvalue. In this paper we use the inverted dominant root of the autoregressive part of the inflationary process of each country as a measure of the degree of inertia displayed by the series.

III. Preliminary Data Analysis

In this section we present a preliminary analysis of the time series properties of monthly inflation rates for sixteen countries. Our main interest is to analyze several measures of persistence using a full sample. In section IV, on the other hand, we analyze the stability through time of our inertia measures, and try to detect structural breaks in the dynamics of the different inflationary processes.

III.1 The Data

We use data from the IMF's International Financial Statistics (IFS) on monthly CPI inflation, for 1970 through 1994, for 16 advanced and developing countries: Argentina, Brazil, Chile, Colombia, France, Germany, India, Israel, Italy, Japan, Mexico, Spain, Sweden, Switzerland, United Kingdom and the United States. In selecting these countries we made an effort cover a wide variety of macroeconomic experiences and of inflationary histories. In Table 1 we present the average, coefficients of variation, and minimum and maximum of the monthly inflation rates. As may be seen, the diversity in the sample is quite remarkable: while Argentina and Brazil had monthly average inflation rates of around 10 percent, Germany and Switzerland had monthly rates in the order of 0.3 percent.

Table 1

Inflation in Several Countries

III.2 Inflationary Persistence: Full Sample Measures

Table 2 presents non parametric variance ratio tests, as well as Dickey-Fuller unit root tests for the sixteen countries in our sample.[8] [9] As may be seen, for all countries the variance ratios clearly decrease as *k* increases. The pace at which the ratio decreases is quite different across countries, however, suggesting that in some of these nations the degree of inflationary persistence may be higher than in others. Moreover, as may be seen from this table (as well as from Figure 1) in some countries --Switzerland, Japan, Spain, for example -- the ratio decreases very rapidly, indicating that inertia has not been an issue. For the period under study This is confirmed by the variance ratio plots (that include confidence intervals) presented in Figure 1.

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[8] In the frequency domain, a non-stationary series will present a large spectrum at low frequency. The variance ratio is an asymptotically equivalent estimate of the spectral density at frequency zero, and therefore we provide estimates of this statistic. In the time space representation a series will be non-stationary if it has a unit root in its auto regressive part. Unit roots test like the Dickey-Fuller, Phillips-Perron and Sargan-Barghawa target that feature of a first order, autoregressive, non-stationary time series. The analysis of the statistical properties of unit root tests is far from being the goal of this paper. However, it is important to keep in mind that most of the unit root tests are linear combinations of the two same statistics and that consequently there is no most powerful test of the unit root hypothesis. In this paper we will use the Dickey-Fuller type of test, by far the most popular among applied economists, in order to identify extreme persistence.

[9] We removed existing monthly additive seasonal components by a standard regression method in all those countries for which the seasonal components are statistically significant. Jaeger and Kunst (1990) show that the method used to seasonally adjust series affects the measured persistence of a series. In particular filtering the data through X-11 methods spuriously introduces persistence into the series. Accordingly we use the simpler regression on dummies method.

It is particularly interesting to contrast Brazil and Mexico, on the one hand, with more developed, low inflation countries like Japan, Switzerland or Sweden. In the former group we can observe that although the variance ratios decrease as k raises, even for k as large as 36 months the variance ratios are still significantly different from zero. In contrast, in the second group of countries, the variance ratios are not significantly different from zero for k's between 18 and 24 months.

Figure 1

Variance-Ratio Tests

Table 2 also shows full sample augmented Dickey-Fuller tests. The table presents the t-statistic for the autocorrelation coefficient obtained in an OLS regression including intercept, deterministic trend and 4 lags of inflation first differences. Using the full sample we are not able to reject the unit-root hypothesis in six of the sixteen countries: France, Israel and Mexico, at 95 percent confidence; Brazil, Chile and Italy at 99 percent confidence.

Table 2

Full Sample Non-Stationarity Tests

From a policy perspective a key question is whether the domestic rate of inflation converges to "world" inflation, and, if it does, the speed at which this convergence takes place. When dealing with convergence we can statistically distinguish three distinct cases: (1) inflation has a unit root and, thus, does not converge to a steady state value; (2)

inflation is white noise, and thus, there is no inflationary inertia; and (3) there is some degree of inertia, with inflation converging slowly to its steady state value. In this latter case we are interested in investigating whether different countries exhibit different speeds of inflationary convergence.

In order to address this issue we proceeded as follows: we fitted MA(1), MA(2), AR(1), AR(2), ARMA(1,1) and ARMA(2,2) for all 16 countries for 1970:01 to 1996:08. In addition, we fitted higher order autoregressive models in all those cases where the correlograms and partial correlograms suggested that these were plausible processes. Table 3 summarizes these results. For each country we indicate whether the series was previously seasonally adjusted, the various models that were fitted to the series, and the Q-statistic white-noise residuals. As may be seen, these results are quite interesting:
First, they suggest that pure moving average representations generally perform poorly in terms of Q-statistics. Second, low order autoregressive process seem to do a good job fitting the inflationary process in some countries. This is the case of Brazil, Colombia, India, Mexico, Sweden, Switzerland, United Kingdom and the United States. In particular, an AR(1)performs well in Brazil, Colombia, India and Mexico. The other four cases are well represented by an AR(2) process. Third, some more developed and relatively low inflation countries such as France, Germany, Italy, Japan and Spain show low Q-statistics in low order ARMA representations.[10] However, these models present the standard near root-cancellation problem. In all 5 cases, a very high inverted AR root has as a counterpart a very high MA inverted root, and therefore these results must be carefully interpreted.[11]

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[10] Inflation in Germany and Italy could also be fairly well represented by an AR(2) process.

[11] To understand this problem, notice that the white noise process

$$y_t = e_t$$

might be rewritten by multiplying both sides of last equation by $(1 - rL)$, and therefore

As an illustration of this problem, consider the case of Germany. The AR(1) and AR(2) representations feature dominant inverted roots of 0.36 and 0.61 respectively. In turn, the MA(1) has a low negative root of -0.29. However, the ARMA(1,1) and ARMA(2,2) representations both present autoregressive and moving average dominant roots close to unity. Finally, the inflation processes of Argentina, Chile and Israel do not seem to be well represented by any of the models presented in the table. This situation may be due either to the fact that we need to specify a higher order model for inflation, or, as we will discuss in next section, to the existence of a structural break in the inflation autoregressive process.

Table 3

Full Sample Measures of Persistence

In summary, with few exceptions, a low order autoregressive process seems to be able to fit the inflationary process observed for most countries in our sample. Interestingly enough the estimated dominant inverted roots differ dramatically across countries, ranging from a 0.09 for Spain, 0.16 for Sweden and 0.38 for Japan, to a high 0.91 for Brazil.[12]

IV. Is the Degree of Persistence Stable Through Time?

The preliminary analysis presented in the preceding section assumes that inflationary persistence is stable through time. This, however, needs not be the case. In fact, since inflation inertia is clearly dependent on institutional price arrangements --

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$$(1-rL)y_t = (1-rL)e_t$$

it is also a valid representation of y_t for any value of r.

12 For this comparisons we consider the highest inverted root of those models with a relatively good Q-statistic, and with no indication of the root cancellation problem.

including indexation rules, mechanisms determining staggered contracts, the degree of credibility of the exchange rate policies and other economic variables --, there are good reasons to believe that inflationary persistence will change through time. In particular, one would expect that a change in the exchange rate regime -- from a crawling peg to a fixed rate, for instance -- will result in a reduction in the degree of persistence.

In this section we investigate whether the data exhibits structural breaks in the degree of persistence. We do this by calculating unit rolling ADF's and sequential Perron tests. The rolling ADF's tests were constructed by rolling a 36-months sample along the full period under analysis. For each one of these sub-samples we performed an ADF test including a time trend. A full-sample interpretation of this test is that if the minimum t-ratio obtained is larger (in absolute value) than the critical value, then the unit root hypothesis can be rejected for the whole sample.[13] The sequential Perron test provides an alternative to this procedure. In this case, a sequence of full sample ADF tests are performed. In each one of them we allow for a structural break in the intercept as well as in the time trend. The break is sequentially moved one period ahead, until we reach the end of the sample. We then compare the minimum t-ratio obtained in the lagged inflation term with the critical value. Table 4 reports the results obtained with these two tests. We reject the full sample unit-root hypothesis with 95% confidence in all cases. The implications are very clear: independently of the autoregressive roots presented in table 3, after allowing for structural breaks and parameter instability, the inflation rate seems to have been stationary for all countries in our sample.

[13] Critical values for this test are provided in Banerjee, Lumsdaine and Stock (1992). The critical value is tipically larger in absolute value than the standard Dickey-Fuller critical value.

Table 4

Unit-Root and Trend Break Hypothesis Tests

In order to investigate whether the degree of inertia has been stable we estimated, for each country, sequential Quandt Likelihood Ratio tests. We construct these tests by performing a sequence of full sample Quandt Likelihood ratio tests using an AR(1) model specification, in which we allow for a change in the autorregressive coefficient. The break point is moved sequentially along the sample. Again, a full sample interpretation of this long series of tests is that if the maximum QLR statistic obtained is larger than the 95 percent critical value, we can not reject the null hypothesis of absence of structural break in the sample.[14] It is important to notice that this procedure is not subject to the standard problem of choosing the break point date through data inspection.

Column 3 of table 4 summarizes the sequential QLR test results. We strongly reject the hypothesis of absence of structural break in the cases of Argentina, Chile, France, Israel, Italy and Mexico. Unfortunately, because we needed at least 50 observations post event in order to perform the test, we were unable of examining the possibility of a structural break after the Real program in Brazil. In any case, these results may explain why, as reported in section III, we were unable to fit a unique ARMA representation for the inflationary processes of Argentina, Chile and Israel.

Figure 2 presents the statistics obtained for each of these three procedures. In those cases where the null hypothesis of absence of structural break is rejected, it is possible to approximately determine the date of the break. For Argentina, a structural break in the

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[14] Again, for a more detailed analysis of this testing procedure and of its asymptotic and empirical distributions see Banerjee, Lumsdaine and Stock (1992).

inflationary process seems to have happened in 1989 at the time of the Alfonsin hyperinflation. Structural breaks are detected in the middle of 1975 in Chile, and in the middle of 1985 in the case of Israel. In section V we provide a detailed analyses of these two cases.

Figure 2

Unit-Root and Trend Break Tests:

Rolling ADF, Sequential Perron and Sequential QLR

IV.1 Univariate Rolling Estimates of Inflationary Inertia

For each country we estimated AR(1) and AR(2) rolling regressions, as a way of further analyzing the changing degree of inertia.[15] Each regression is estimated for a rolling 36-months sample. Figure 3 plots the rolling autoregressive coefficients obtained in the AR(1) regressions, with two-standard deviation bands. The largest inverted autorregressive roots of the AR(2) rolling estimations are plotted, for each country, in figures 4.[16] [17]

The rolling AR(1) estimates reported in figure 3 are quite revealing. First, they show an important diversity of country experiences. While in some nations the degree of inertia is low and stable -- Sweden, Switzerland, for example --, in others it is highly volatile during the period under analysis (Israel, for example). Second, they clearly suggest

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[15] Alternative ways of facing the parameter instability problem are recursive regressions and Kalman filter. In both of these methodologies, the estimate at each date considers all the information contained in the sample up to that point. The rolling regression approach is more prone to capture changes in the inertia regime because all the information used for the estimation is renewed at short intervals.

[16] When the largest root is imaginary, we use the modulus of the root.

[17] In the figures, we use as a convention to date the autoregressive coefficient obtained in each rolling regression in the last month included in the sample. Montecarlo exercises showed that using rolling regressions with 36 observations, a structural break is statistically detected when at least 20 months of the new regime are included in the sample. Hence, changes in the rolling coefficients should be associated with

that, in many countries, the degree of inflationary inertia varies significantly through time. This is, for instance the case of Argentina, which exhibits no significant inflationary persistence in the early 1970s, followed by a significantly increase in inertia during the 1980s, and a marked decline in the 1990s. This latter development is related to the adoption of the convertibility stabilization plan in 1991. The estimates for Italy also suggest a rapid significant change in the degree of inertia. This is consistent with the results on the Quandt rolling structural break test reported above. Third, a number of countries seem to have lost (at least temporarily) their nominal macroeconomic anchor. In this case the AR coefficient is not significantly different from unity at conventional levels. This appears to be the case in Argentina in 1989, Brazil during the late 1980s and early 1990s. Fourth, some countries appear to have been able to defeat inflationary persistence, with the estimated coefficients of inertia collapsing from a positive – and in some cases very high – value to one that is not significantly different from zero. An important question is whether those countries that have succeeded in eliminating inertia have followed similar policy courses. In section V we address this issue from the point of view of nominal exchange rate policy, and analyze in greater detail the evolution of inertia in three countries that, in the last ten years or so, have implemented nominal exchange rate anchor based stabilization programs.

Figure 3

Rolling AR(1)

Figure 4

Rolling AR(2)

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events happening ona and a half years before the date in the figure.

For most countries the results obtained from the estimation of rolling AR(2), and reported in Figure 4, tell a very similar story from those presented in Figure 3 for AR(1). In the following cases, however, there are some differences: Colombia, France, Israel, Japan, Spain and Sweden. However, as reported in Table 3, the AR(2) representation appears significantly preferable to the AR(1) only for Japan, Spain and Sweden. In the other three cases, there does not seem to make a difference, at least for the full sample**.**

The case of the United States deserves particular attention. As may be seen, figures 3 and 4 suggest strongly that the degree of inflationary persistence has not been stable through time. Between 1970 and 1975 the estimated AR coefficient it was not significantly different from zero, suggesting that inflation was characterized by white noise and that there was an absence of inflationary inertia. Starting in 1976, however, there is a marked increase in the degree of persistence. This peaked in the early 1980s, when the point estimate of the AR coefficient was 0.8, not very different from that of the highly inflationary nations of South America. From then onward, inflationary inertia has experienced a somewhat gradual decline, until once again experiencing, in the mid 1990's a virtual disappearance. These results suggest that analyses that assume a stable degree of persistence, such as that of Fuhrer and Moore (1995), may result in highly misleading policy implications.

IV.2 Multivariate-Based Measures of Inflationary Inertia

A possible limitation of the results reported above is that they have been obtained on the bases of univariate time series techniques, without controlling for other variables. However, the omission of possibly important determinants of inflation may introduce some biases in the estimation. As a way of dealing with this potential source of bias we estimated the same set of AR(1) rolling regressions for some of the countries in our sample including as an additional explanatory variable the rate of growth of nominal money (IFS line 34).[18] Figures 5 show the results obtained in this case for four countries: Argentina, Chile, Israel and Mexico. It is clear that in all cases, the differences with respect to the results of figures 3 are unimportant. In the cases of Chile and Mexico we went a step further, and reestimated the AR(1) regressions controlling by estimated excess supply for money. This latter variable was obtained, in turn, as residuals from error correction based estimates of the demand for money. As may be seen from this figure, the broad results reported above are affected by the inclusion of additional variables into the analysis.[19]

Figure 5

Rolling AR(1) including money growth

IV.3 Inflation Levels and Persistence

The results summarized in Figures 3, 4 and 5 present a simple measure of inflationary persistence for an heterogeneous set of countries. A key contribution of this paper is that we have allowed this measure of inertia to vary over time. This exercise provides a rich panel of inflationary persistence data, that contains both cross-country and time series variation. An interesting question – and one that has important policy implications -- is whether countries and periods of time where inflation was high, tended also to present high inertia. Figure 6 presents scattered diagrams of average inflationary inertia -- measured as the average autoregressive coefficient obtained in the 12 rolling

[18] There is no monthly measure of money for several of the countries. For others the series are too short or too incomplete.

[19] We also estimated impulse response functions for laternative subsamples. Although we were unable to make fine timing distinctions, the estimates provided support to the results reported above.

regressions for each year – and average annual inflation for different subsets of the countries in the sample. Interestingly enough, these figures show that there is a positive relationship between the average level of inflation and the measure degree of inertia. As can be seen from the different panels in this Figure, this relationship is present at different levels of average inflation. In order to show that this graphical evidence is also statistically significant we performed panel regressions of inflation on persistence. We used the same information plotted in Figure 6 to run panel regressions using pooled least squares, fixed effects and random effects estimators. We also divided the sample in four different sub-samples according to the average level of inflation. Table 5 shows the results obtained. As may be seen, these results confirm those presented in Figure 6 and discussed above. In summary, all the regressions show a positive and statistically significant relationship between inflationary persistence and inflation levels. The results are robust to the estimation procedure and the sub-sample selected.

A possible explanation for this finding is that as the inflation level increases society's develop institutional arrangements to protect themselves from the effects of inflation. Many of these protective measures – and more specifically, indexation mechanisms – will tend to increase the degree of inflationary inertia, by adjusting prices according to (some measure) of past inflation. In order to find out whether countries with extensive indexation practices indeed exhibit higher inertia, we performed some preliminary non parametric analysis. These results indicate that it is possible to reject the hypothesis that high indexation and no indexation countries come from the same sample. Moreover, average inflation rates in high indexation countries are significantly higher than in countries where there are no generalized indexation practices.[20]

Figure 6

Inflation and Persistence

Table 5

Persistence and Inflation: Regression Results

V. Inertia and Inflation Convergence: Three Cases

An important goal of most stabilization programs in the developing (or emerging) economies is to reduce the domestic rate of inflation to levels comparable to those of "world inflation." In other words, stabilization programs are usually aimed at producing inflationary convergence. The speed at which inflation is reduced is important, however. Under fixed nominal exchange rates a slow inflationary convergence will result in real exchange rate overvaluation, speculation and, in the end, can even generate a crisis. If inertia is eliminated, or greatly reduced, inflationary convergence will be achieved rapidly; on the contrary, if inertia is very high, convergence will tend to take a long time.[21] Historically many countries – including a number in our sample -- have attempted to achieve inflation convergence by implementing stabilization programs based on the adoption of a rigid (and in some cases totally fixed) nominal exchange rate. In this subsection we analyze briefly the experiences of three countries in the sample, that have adopted nominal exchange rate anchors stabilization programs in the last twenty years or

[20] These preliminary results are available from the authors on request.

[21] Naturally, in making these statements we are assuming that the fundamental macroeconomic stance is

so: Chile, Israel and Mexico. Our main interest is to analyze, in the light of the battery of tests presented in the preceding sections, whether the adoption of this type of program indeed resulted in a reduction (or elimination) of the degree of inertia. We begin our discussion with a direct comparison of the Chilean and Mexican experiences of the 1970s and 1980s. We then move to the Israeli case.

Chile (1978-82) and Mexico (1987-94)

The Chilean reform program was initiated in 1975, ten years prior to the launching of the Mexican reforms. Both programs, however, shared a number of features: (a) drastic opening of the economy; (b) ambitious privatization and deregulation; and (c) a stabilization program based on a predetermined, nominal exchange-rate anchor, supported by (largely) restrictive fiscal and monetary policies. Both programs were rooted in the notion that a predetermined rate of nominal devaluation -- including a fully fixed nominal exchange rate -- would constrain price increases, reduce inflationary expectations and ensure that monetary and fiscal authorities would behave conservatively. In Chile, the exchange-rate-based stabilization program had two phases: from February 1978 through June 1979 the rate of devaluation was preannounced and deliberately set below the ongoing rate of inflation. In June of 1979 the nominal exchange rate was pegged at 39 pesos per dollar. The authorities announced that this new exchange rate would be in effect for the "indefinite" future.[22]

The Mexican program followed a somewhat different pattern. Between February and December of 1988 the nominal exchange rate was fixed, becoming the fundamental

conducive to inflationary reduction.

22 Parts of this comparison draws on Edwards (1997).

anchor of the anti-inflationary effort.[23] Between January 1989 and November 1991 a system based on a pre-announced rate of devaluation was in effect. As in Chile a decade earlier, during this period the actual rate of devaluation was deliberately set below the ongoing rate of inflation as a way of reducing expectations and price increases. The amount by which the peso was devalued was successively reduced from one peso per day in 1989, to 80 cents in 1990, to 40 cents in 1991 and to 20 cents in 1991. In November 1991 the authorities added some flexibility to the system, and an exchange rate band with a sliding ceiling and a flat floor was adopted. This measure was justified on two grounds: (a) it was supposed to discourage short term capital inflows; and (b) it was supposed to allow some real exchange rate corrections. Until October 1993 -- when the NAFTA controversy heated up in the United States -- the actual peso/dollar rate was extremely stable, remaining in the lower half of the band. During 1994, however, and as a result of political and other developments, the exchange rate came under considerable pressure, moving towards the top of the band. As is well known, the attempt to widen the band in late December failed badly. It was too little, too late.

From early on, Chilean and Mexican authorities understood that consistent fiscal and monetary policies were a precondition for the success of the exchange-rate-based stabilization programs. In fact, in both countries the primary balance of the public sector was under control even before the adoption of the exchange-rate anchor, and monetary policy was largely restrained during most of the period.[24]

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[23] During October-December 1987, the first months of the *Pacto*, nominal wages provided the anchor to the system. According to Vela (1993) the move to an exchange-rate anchor in February 1988 was, in part, the result of labor union pressure.

[24] During the early years of the Mexican program, monetary policy was guided by a dual objective: on the one hand the authorities were interested in reducing interest rates -- which had reached extremely high real levels in 1987 and 1988 --, and on the other hand they wanted to make sure that domestic credit policy would be consistent with the predetermined nominal exchange rate (Aspe 1993, p. 37).

The architects of both the Chilean and Mexican programs expected that a predetermined exchange-rate would affect inflation through, at least, two channels. First, in an open economy a fixed exchange rate would impose a ceiling to tradable inflation; second, the new exchange rate policy was expected to generate a major break in inflationary expectations and in inflationary inertia. Mexico's finance secretary Pedro Aspe put it the following way: "[Pegging the exchange rate] and...lowering... trade barriers in the tradable sector were indispensable for breaking down inertia. Consensus and a tendency toward purchasing power parity could reinforce each other to bring down inflation" (Aspe 1993, p. 44).

In spite of these similarities, there were some important differences between the Chilean and Mexican programs. Perhaps the most important one referred to wage rate policy. Beginning in 1976, Chile adopted a formal backward looking wage indexation scheme, through which all wages in the formal sector were periodically raised to compensate for the cumulative rate of inflation of the previous period (Edwards and Edwards 1991, Ch. 6). In contrast, the Mexican *Pacto* encompassed a social and economic agreement among the government, the private sector and labor unions aimed at establishing guidelines for price, wage, and exchange rate changes.

In both countries inflation was reduced, but at a significantly slower pace than what the authorities had anticipated. As their reforms proceeded and became consolidated, both countries were subject to very large capital inflows which helped finance increasingly large current account deficits and generated very large real exchange rate overvaluation (Edwards 1996). In Chile the current account deficit exceeded 12% of GDP in 1981, while in Mexico it surpassed 7% of GDP both in 1992 and 1993. As it turned out, these very large deficits proved to be unsustainable and both countries ended up facing major crises: the Chilean crisis erupted in June of 1982; the Mexican crisis in December 1994.

How successful were these two nominal exchange rate-based stabilization programs in reducing (or eliminating) inflationary inertia? We can address this issue with the aid of the statistical analysis presented in this paper. In particular, we can use the different estimates of inertia developed in the preceding section to analyze whether, as the architects of the programs had intended, the degree of inflationary persistence declined after the exchange rate anchor program was enacted. In Figure 7 we present once again our estimates of inertia for these two countries using the AR(1) and AR(2) procedures.

Figure 7

Inflationary Inertia: Chile and Mexico

The results are quite interesting. In the case of Chile, by the end of 1979 there is a temporary decline in inertia – both the AR(1) coefficient and the dominant root in the AR(2)-based measure decline temporarily to 0.33. However, from then onward measured inertia started to increase once again, reaching more than 0.8 in the months immediately preceding the abandonment of the fixed exchange rate in June 1982. There are two possible explanations for this behavior. First, due to the mandated backward-looking wage indexation mechanism, structural inertia remained in the system in spite of the adoption of the anchor. Second, and as argued in Edwards (1998), as real exchange rate devaluation became more and more acute the degree of credibility of the program was eroded, with economic agents incorporating an expected change in the exchange rate regime, and a resumption of the crawling peg policy, in their pricing decisions.

As in Chile, Mexico experienced an initial decline in persistence. After a few months, however, measured inertia had once again increased (Figure 7), and in mid 1994, seven years into the program, it was still in the range of 0.6 to 0.7. This high degree of inertia contributed to the creation of an acute real exchange overvaluation, which contributed to the unleashing of the Mexican currency crisis of December 1994 (Dornbusch and Warner 1994).

Israel

During 1983-84 Israeli inflation averaged almost 300% per annum in year. This was the result of serious macroeconomic imbalances, including double digit fiscal deficits and an entrenched and perverse inflationary expectations process. In mid-1985 Israel launched an ambitious stabilization program based on three important elements: (1) a control of the fiscal deficit; (2) a social pact agreed among representatives of labor, government and the private sector; and (3) a fixed (but adjustable) nominal exchange rate. As in the cases of Chile and Mexico discussed above, the exchange rate anchor policy was aimed at breaking expectations, placing a ceiling on tradables inflation and reducing inertia. In many ways the program was a success: in 1986 inflation was 20%, down from 185% in the previous year, and by 1987 inflation had been further reduced to 16%.

As in Chile and Mexico, this program generated a process of real exchange rate appreciation which hurt the degree of international competitiveness of Israeli exports. Between 1985 and 1988 this problem was handled through periodic nominal devaluations, that were followed by a new peg. In January 1989 , however, a new exchange rate policy based on a crawling exchange rate band was implemented as a way to combine the dual objectives of defeating inflation and avoiding an exchange rate crisis.

Figure 8 presents the evolution of Israel's nominal exchange rate from mid 1985 until December 1993. Initially this band allowed the exchange rate to fluctuate by ± 3% around a fixed central parity. Subsequently the width of the band was increased to ± 5%, and in December 1991 the sliding band was replaced by a crawling band characterized by a moving central parity based on **expected** inflation rate differentials.

Figure 8

Israel's Exchange Rate Band

The move from the sliding to the forward looking crawling band in late 1991, responded to two factors. First, the sliding band system was generating significant uncertainty. Every time the public expected that a realignment of the central parity was about to take place, there was increased speculation and domestic interest rates jumped up. Second, by basing the rate of the crawl of the band on expected, rather than lagged inflation, the authorities expected that they would affect inflationary expectations and would reduce -- if not eliminate -- inflationary inertia. With the exception of the first band that operated from January through June of 1989, the band width was maintained at ± 5%.

The adoption of the forward looking band was, in many respects, successful. First, the speculation surrounding the realignment of the central parity was greatly reduced, and interest rates declined and became less volatile. Second exchange rate overvaluation was avoided, and exports performance improved significantly. Third, in the early 1990s inflationary inertia was greatly reduced. Fourth, as predicted by the program inflation declined gradually through 1995. And fifth, this band allowed the Israeli economy to accommodate in a non-traumatic fashion changes in fundamentals, including episodes of large capital outflows -- in the last quarter of 1992 capital outflows, for example, amounted to over 20% of international reserves.

Figure 9

Inflationary Inertia: Israel

From the perspective of this paper, an important question is how the Israeli exchange rate-based program affected inflationary inertia. This can be addressed by reexamining the inertia measures computed above, which are reproduced in Figure 9 for convenience. This figure is highly revealing and provides an interesting contrast with those for Chile and Mexico displayed above. As may be seen, the preferred AR(1) measure of inertia very clearly indicates that already by late 1986 inflationary persistence in Israel was experiencing a rapid collapse. In fact, it is not possible to reject, at conventional significance levels, the hypothesis of complete absence of inertia between late 1988 and mid 1990. These estimates, however, quite strongly suggest that by 1991 the degree of inertia was once again on the rise, and that by 1992-93 it was almost as high as it had been in 1983-86. The most likely explanation for this result is that by then – with inflation in the 20% level for the sixth year in a row -- it had become clear that the war against inflation had run into a standstill. The program had run into credibility problems, as old indexation practices had slowly creeped back into the system.

VI. Concluding Remarks

This paper has empirically analyzed the extent of inflationary inertia in a set of countries. A contribution of the paper has been to assume that the degree of inflationary persistence can change through time, and to associate changes in the time varying measures of inflationary inertia to specific economic events for a sub-set of countries. The paper also discusses preliminary evidence regarding the effect of indexation practices and credibility on inflationary persistence.

We analyzed the time series properties of monthly inflation data from 1970 through the end of 1996. We have performed a battery of tests: (a) Rolling Augmented Dickey-Fuller tests and (b) Sequential Perron tests to test the unit-root hypothesis in the presence of structural breaks, and (c) Sequential Quandt likelihood ratio tests to analyze the instability of the autoregressive coefficients. We found that allowing for structural breaks, it is possible to reject the unit root hypothesis for all countries in our sample, and that it is possible to reject the hypothesis of absence of structural break in the first order autoregressive coefficient in Argentina, Chile, France, Israel, Italy and Spain. In addition, the reported estimates of the rolling AR(1) coefficients and the dominant inverted roots of the rolling AR(2) estimates, clearly show that inflationary inertia greatly varies both across countries and across different periods of time within a country. We have shown that there exists a clear connection between our estimates of inertia and the levels of inflation in the countries in our sample. Episodes (time and place) characterized by a high degree of inflationary persistence also feature a high level of inflation.

We ended the paper discussing in more detail the evolution of inflationary inertia in three countries that have relied, at one point or another, on a nominal exchange rate anchor to achieve price stability: Chile, Israel and Mexico. We have shown that the measure of inertia falls at the beginning of these stabilization programs, but interestingly enough, rises back as time goes by and the programs fail to cut inflation down to “world” levels. Unfortunately, due to data constraints we were not able in analyzing the last Argentinean stabilization program.

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Table 1. Inflation in Several Countries *

	Average Rate of Inflation	Coefficient of Variation	Min. Monthly Inflation Rate	Max. Monthly Inflation Rate
Argentina	9.3	1.8	-6.0	196.6
Brazil	10.2	1.2	-0.3	80.7
Chile	3.7	1.8	-8.1	87.5
Colombia	1.7	0.6	-2.0	6.9
France	0.5	0.7	-0.3	1.9
Germany	0.3	1.1	-0.5	1.8
India	0.7	1.5	-2.9	4.3
Israel	3.5	1.3	-1.4	27.5
Italy	0.8	0.7	-0.4	3.1
Japan	0.4	2.0	-1.1	4.1
Mexico	2.5	1.0	0.0	15.5
Spain	0.8	1.1	-4.4	9.0
Sweden	0.6	1.1	-0.5	3.3
Switzerland	0.3	1.3	-0.7	2.1
U.K	0.7	1.1	-0.9	4.3
U.S.A.	0.4	0.7	-0.5	1.8

* Percentage, except Coeff. of Variation. Sample from 1970:01 to 1996:08.

Figure 1. Variance Ratio Test

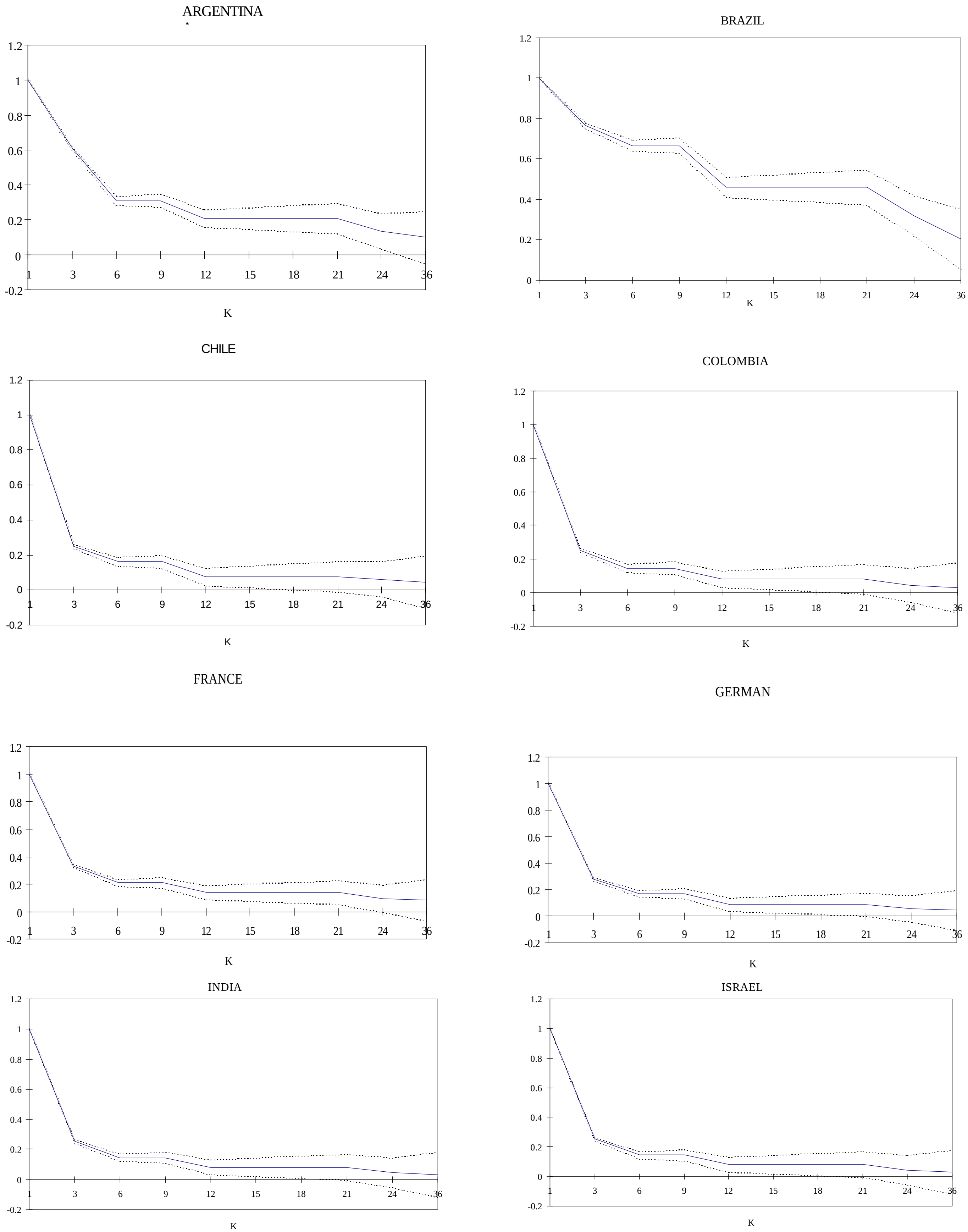


Figure 1. Variance Ratio Test (cont).

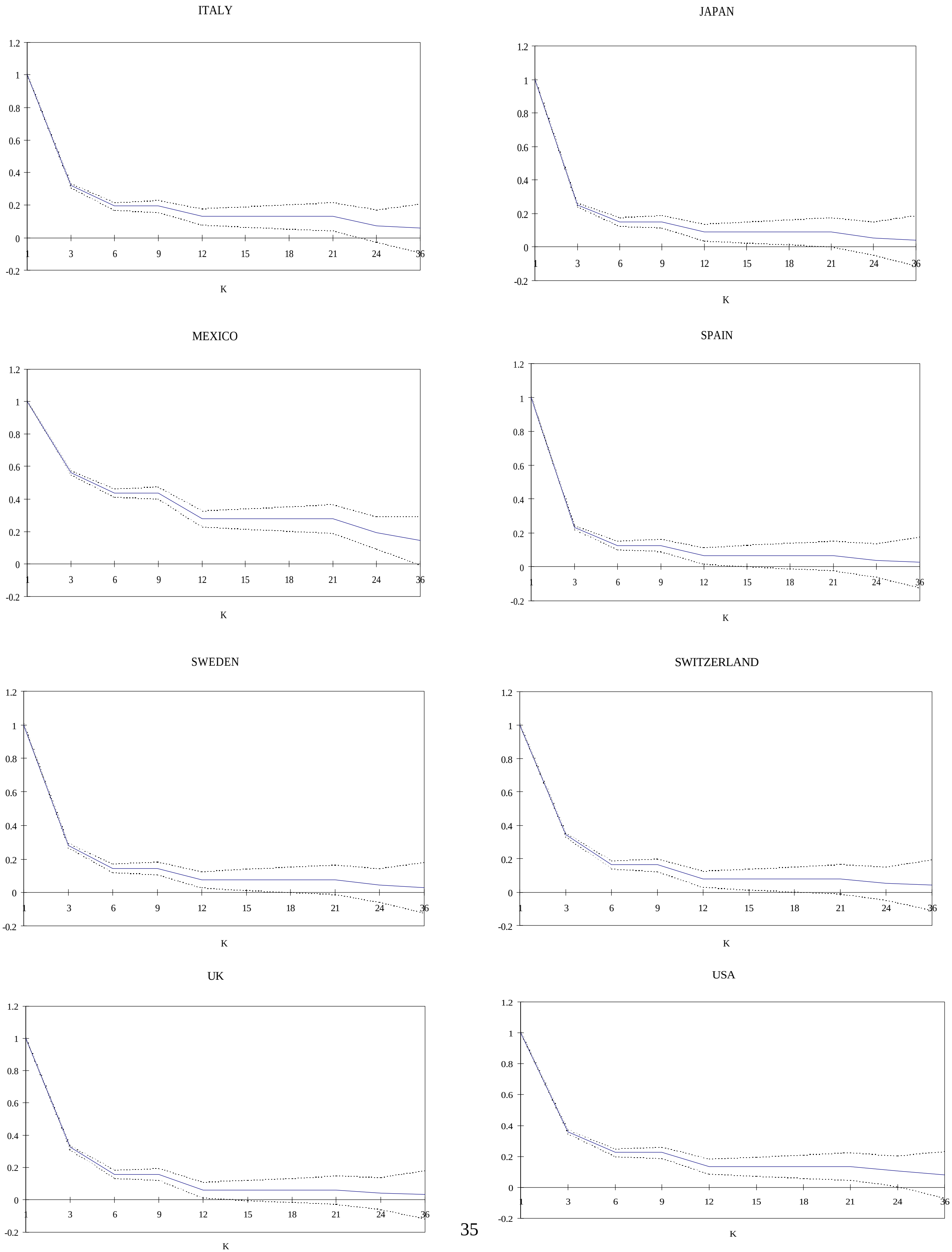


Table 2. Full Sample Non-Stationarity Tests

Country	Seasonal Adjustment	Variance Ratio Test					ADF *
		k=3	k=6	k=12	k=24	k=36	
Argentina	NO	0.61	0.31	0.21	0.14	0.10	-4.80
Brazil	NO	0.76	0.67	0.46	0.32	0.20	-3.89
Chile	NO	0.25	0.16	0.08	0.06	0.05	-3.78
Colombia	YES	0.25	0.15	0.08	0.04	0.03	-6.76
France	YES	0.34	0.21	0.14	0.10	0.08	-3.40
Germany	YES	0.28	0.17	0.09	0.06	0.05	-5.54
India	YES	0.25	0.15	0.08	0.04	0.03	-4.16
Israel	NO	0.25	0.15	0.08	0.04	0.03	-2.57
Italy	YES	0.32	0.19	0.13	0.07	0.06	-3.87
Japan	YES	0.25	0.15	0.09	0.05	0.04	-5.34
Mexico	NO	0.56	0.44	0.28	0.19	0.14	-3.28
Spain	YES	0.23	0.13	0.07	0.04	0.03	-6.13
Sweden	YES	0.28	0.15	0.08	0.04	0.03	-6.85
Switzerland	YES	0.34	0.16	0.08	0.05	0.04	-6.48
UK	NO	0.33	0.16	0.06	0.04	0.03	-5.89
USA	YES	0.36	0.23	0.14	0.11	0.08	-4.02

* Critical Values are -3.42 (5%), -3.98 (1%)

Table 3. Full Sample Measures of Persistence

Country	Seas. Adj.	Model Specification	Inverted AR Roots				Inv. MA Roots		Q-Statistic p=5*	p=10**
Argentina	NO	AR(1)	0.68						15.6	62.4
		AR(2)	0.65	0.05					12.8	61.0
		MA(1)					-0.35		18.7	31.4
		MA(2)					-0.18+0.26i	-0.18-0.26i	13.3	23.1
		ARMA(1,1)	0.66				-0.04		13.4	61.3
		ARMA(2,2)	0.366-0.27i	0.36+0.27i			0.02+0.54i	0.02-0.54i	2.2	48.9
Brazil	NO	AR(1)	0.91						9.4	13.3
		AR(2)	0.90	0.06					8.1	11.5
		MA(1)					-0.72		491.1	739.5
		MA(2)					-0.45+0.63i	-0.45-0.63i	259.6	397.8
		ARMA(1,1)	0.90				-0.05		8.3	11.8
		ARMA(2,2)	0.88	-0.48			-0.29+0.29i	-0.29-0.29i	2.3	6.2
Chile	NO	AR(1)	0.46						49.1	92.1
		AR(2)	0.71	-0.37					39.3	69.4
		AR(3)	0.85	-0.29+0.49i	-0.29-0.49i				28.6	46.0
		AR(4)	0.92	-0.05+0.70i	-0.05-0.70i	-0.64			4.2	13.3
		MA(1)					-0.33		137.7	256.6
		MA(2)					-0.17-0.4i	-0.17+0.4i	82.4	159.1
		ARMA(1,1)	0.98				0.83		9.2	26.0
		ARMA(2,2)	0.98	-0.97			0.83	-0.99	9.3	25.2
Colombia	YES	AR(1)	0.49						2.9	7.2
		AR(2)	0.54	-0.09					2.2	6.4
		MA(1)					-0.42		21.4	34.1
		MA(2)					-0.23-0.43i	-0.23+0.43i	8.7	18.0
		ARMA(1,1)	0.57				0.11		2.5	6.7
		ARMA(2,2)	0.52	-0.27			-0.01	-0.20	2.1	6.4
France	YES	AR(1)	0.81						27.8	44.0
		AR(2)	0.89	-0.24					24.6	35.4
		AR(3)	0.95	-0.19+0.52i	-0.19-0.52i				11.7	15.1
		MA(1)					-0.69		327.8	657.0
		MA(2)					-0.4-0.37i	0.4+0.37i	182.5	373.3
		ARMA(1,1)	0.99				0.69		19.4	20.8
		ARMA(2,2)	0.99	-0.22			0.79	-0.53	4.0	5.8
Germany	YES	AR(1)	0.36						20.0	43.1
		AR(2)	0.61	-0.30					8.0	18.9
		MA(1)					-0.29		47.8	93.7
		MA(2)					-0.14+0.38i	-0.14-0.38i	27.9	59.2
		ARMA(1,1)	0.97				0.85		3.9	6.1
		ARMA(2,2)	0.97	-0.41			0.86	-0.50	1.2	3.1

* Critical values at 95% confidence are 9.49 for the AR(1) and MA(1) models, 7.82 for the AR(2), MA(2) and ARMA(1,1) models; 5.99 for the AR(3) model; and 3.84 for the AR(4) and ARMA(2,2) models.

** Critical values at 95% confidence are 16.92 for the AR(1) and MA(1) models, 15.51 for the AR(2), MA(2) and ARMA(1,1) models; 14.07 for the AR(3) model; and 12.59 for the AR(4) and ARMA(2,2) models.

Table 3. (continued)

Country	Seas. Adj.	Model Specification	Inverted AR roots			Inverted MA Roots		Q-Statistic p=5	p=10
India	YES	AR(1)	0.50					19.3	28.3
		AR(2)	0.63	-0.19				16.7	23.5
		MA(1)				-0.42		70.2	96.1
		MA(2)				-0.22+0.32i	-0.22-0.32i	43.1	61.1
		ARMA(1,1)	0.88			0.57		10.7	14.0
		ARMA(2,2)	0.90	-0.42		0.65	-0.57	3.5	6.1
Israel	NO	AR(1)	0.75					45.6	71.6
		AR(2)	0.83	-0.21				36.5	57.5
		AR(3)	0.92	-0.18+0.57i	-0.18-0.57i			23.1	31.1
		MA(1)				-0.71		227.5	451.6
		MA(2)				0.39-0.06i	0.39+0.06i	158.1	302.6
		ARMA(1,1)	0.97			0.64		24.8	36.6
		ARMA(2,2)	0.98	-0.43		0.75	-0.75	99.3	188.2
Italy	YES	AR(1)	0.67					30.7	54.4
		AR(2)	0.85	-0.40				6.7	17.2
		MA(1)				-0.45		230.9	437.8
		MA(2)				0.25-0.54i	0.25+0.54i	102.6	220.1
		ARMA(1,1)	0.97			0.69		4.8	8.1
		ARMA(2,2)	0.97	-0.85		0.70	-0.88	49.4	8.1
Japan	YES	AR(1)	0.38					38.5	86.5
		AR(2)	0.67	-0.38				12.5	35.3
		AR(3)	0.81	-0.29+0.43i	-0.29-0.43i			4.3	19.9
		MA(1)				-0.29		79.0	172.7
		MA(2)				-0.13+0.45i	-0.13-0.45i	46.4	112.7
		ARMA(1,1)	0.98			0.84		1.9	12.3
		ARMA(2,2)	0.97	0.16		0.82	0.18	2.0	12.2
Mexico	NO	AR(1)	0.86					6.5	9.5
		AR(2)	0.87	-0.04				7.0	10.0
		MA(1)				-0.71		386.0	624.6
		MA(2)				-0.45+0.52i	-0.45-0.52i	170.5	280.6
		ARMA(1,1)	0.88			0.08		7.6	10.5
		ARMA(2,2)	0.86	-0.96		0.04	-0.99	6.5	9.2
Spain	YES	AR(1)	0.09					44.8	111.1
		AR(2)	0.50	-0.44				18.0	49.5
		MA(1)				-0.06		50.5	124.5
		MA(2)				-0.02+0.39i	-0.02-0.39i	29.7	79.2
		ARMA(1,1)	0.99			0.94		13.3	20.0
		ARMA(2,2)	0.99	-0.07		0.91	0.13	4.8	13.8

* Critical values at 95% confidence are 9.49 for the AR(1) and MA(1) models, 7.82 for the AR(2), MA(2) and ARMA(1,1) models; 5.99 for the AR(3) model; and 3.84 for the AR(4) and ARMA(2,2) models.

** Critical values at 95% confidence are 16.92 for the AR(1) and MA(1) models, 15.51 for the AR(2), MA(2) and ARMA(1,1) models; 14.07 for the AR(3) model; and 12.59 for the AR(4) and ARMA(2,2) models.

Table 3. (continued)

Country	Seas. Adj.	Model Specification	Inverted AR roots		Inverted MA Roots		Q-Statistic p=5	Q-Statistic p=10
Sweden	YES	AR(1)	0.16				10.4	39.1
		AR(2)	0.44	-0.30			2.3	24.7
		MA(1)			-0.13		13.7	47.8
		MA(2)			-0.07+0.36i	-0.07+0.36i	4.5	29.5
		ARMA(1,1)	0.99		0.93		4.0	14.6
		ARMA(2,2)	0.99	-0.99	0.93	-0.98	3.5	15.8
Switzerland	YES	AR(1)	0.36				15.8	44.4
		AR(2)	0.62	-0.33			1.2	20.0
		MA(1)			-0.25		38.6	92.9
		MA(2)			-0.14+0.48i	-0.14-0.48i	5.9	38.4
		ARMA(1,1)	0.95		0.82		23.3	37.3
		ARMA(2,2)	0.67	-0.97	0.30	-0.93	0.9	15.2
UK	NO	AR(1)	0.45				8.7	42.9
		AR(2)	0.62	-0.24			3.1	34.7
		MA(1)			-0.36		45.6	103.2
		MA(2)			-0.18+0.37i	-0.18-0.37i	23.4	73.4
		ARMA(1,1)	0.96		0.79		14.3	36.8
		ARMA(2,2)	0.98	0.15	0.89	-0.07	2.4	21.5
USA	YES	AR(1)	0.67				23.7	45.7
		AR(2)	0.82	-0.33			5.4	18.1
		MA(1)			-0.48		183.3	350.6
		MA(2)			-0.23+0.55i	-0.23-0.55i	86.6	192.4
		ARMA(1,1)	0.94		0.58		14.6	24.6
		ARMA(2,2)	0.97	0.41	0.81	0.15	5.4	13.5

* Critical values at 95% confidence are 9.49 for the AR(1) and MA(1) models, 7.82 for the AR(2), MA(2) and ARMA(1,1) models; 5.99 for the AR(3) model; and 3.84 for the AR(4) and ARMA(2,2) models.

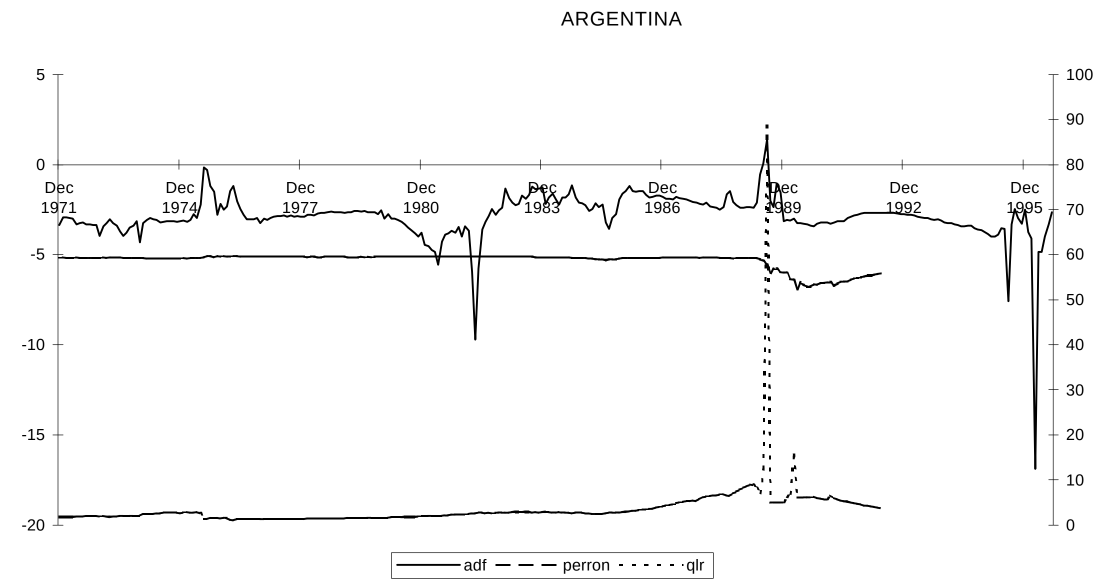
** Critical values at 95% confidence are 16.92 for the AR(1) and MA(1) models, 15.51 for the AR(2), MA(2) and ARMA(1,1) models; 14.07 for the AR(3) model; and 12.59 for the AR(4) and ARMA(2,2) models.

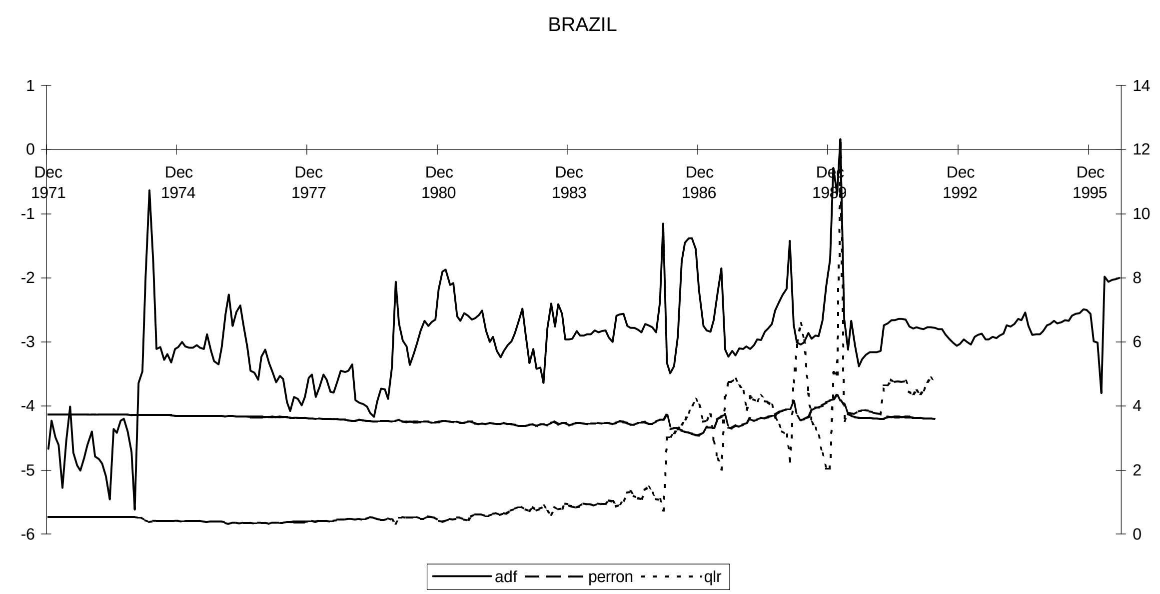
Table 4. Unit-Root and Trend Break Hypothesis Tests

	Rolling ADF	Sequential Perron Test	Sequential Quandt LR
Critical Value (95%)	-4.85	-4.39	27.87
Argentina	-16.89 *	-6.91 *	88.63 **
Brazil	-5.62 *	-4.45 *	12.19
Chile	-8.45 *	-5.52 *	32.10 **
Colombia	-7.56 *	-9.82 *	4.81
France	-6.44 *	-6.16 *	35.06 **
Germany	-6.63 *	-8.27 *	6.79
India	-6.20 *	-9.50 *	4.62
Israel	-19.07 *	-7.85 *	47.20 **
Italy	-6.38 *	-5.95 *	49.49 **
Japan	-7.80 *	-8.33 *	10.55
Mexico	-9.26 *	-7.12 *	18.81
Spain	-8.39 *	-7.58 *	38.62 **
Sweden	-6.65 *	-8.37 *	12.37
Switzerland	-6.27 *	-8.68 *	9.37
UK	-5.53 *	-7.40 *	17.81
USA	-5.08 *	-6.95 *	20.11

* Reject Hypothesis of Unit Root at 95% of Confidence.
** Reject Hypothesis of Non-Structural Break.

Figure 2. Unit Root and Trend Break: Rolling ADF, Sequential Perron and Sequential QLR





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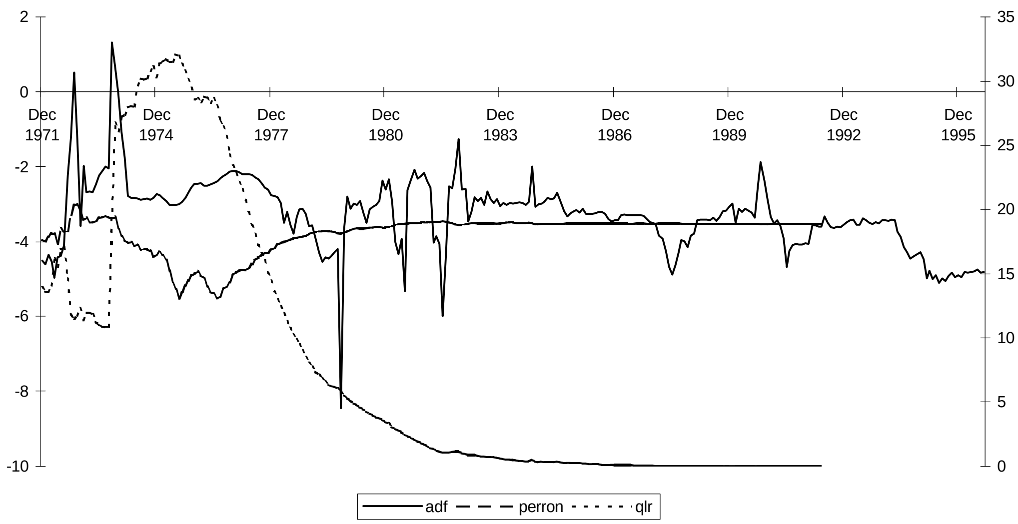

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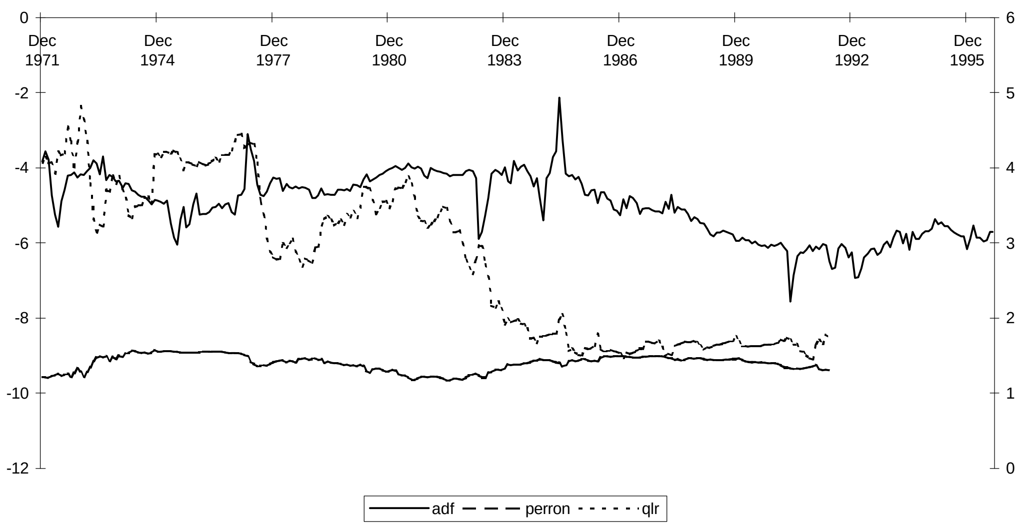

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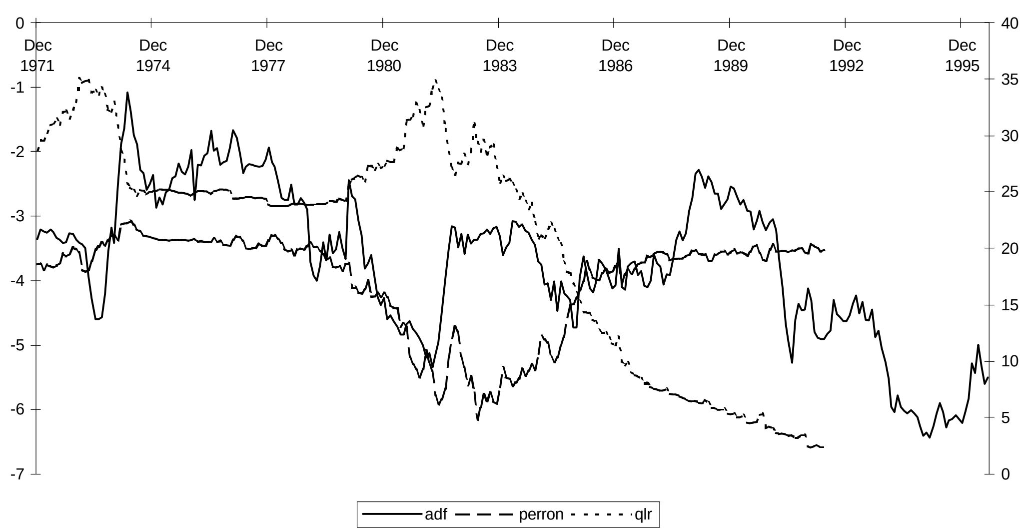

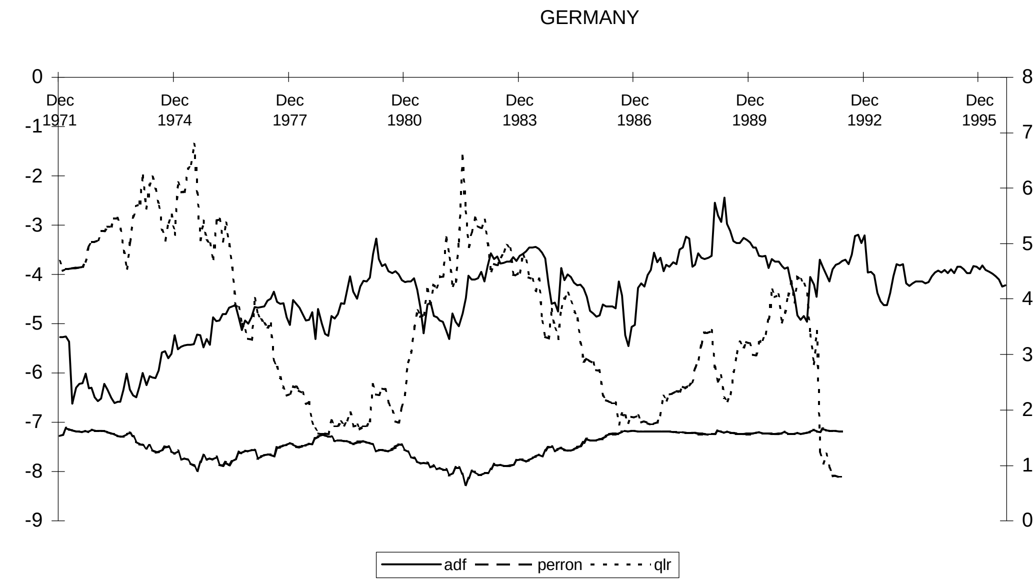
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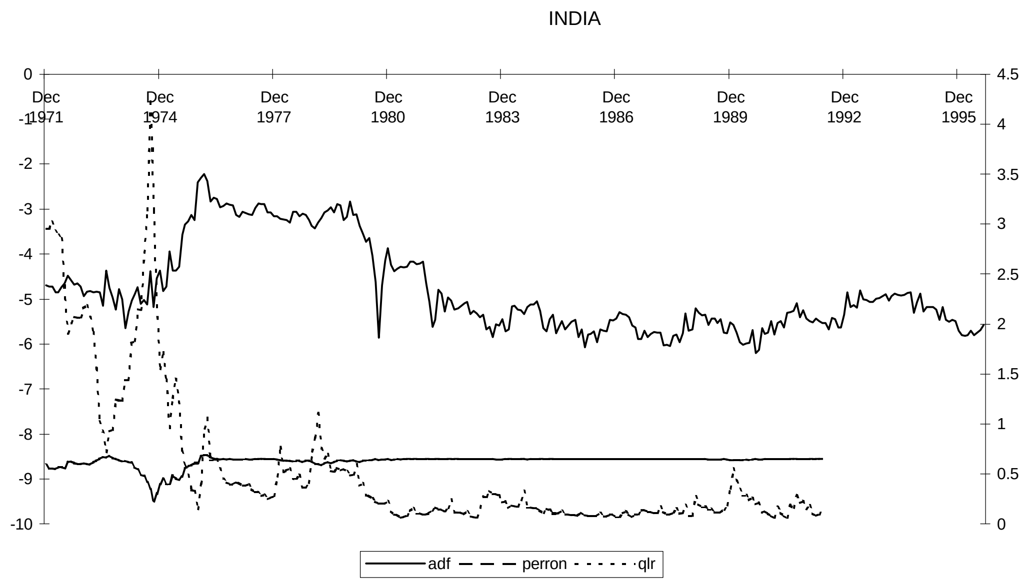
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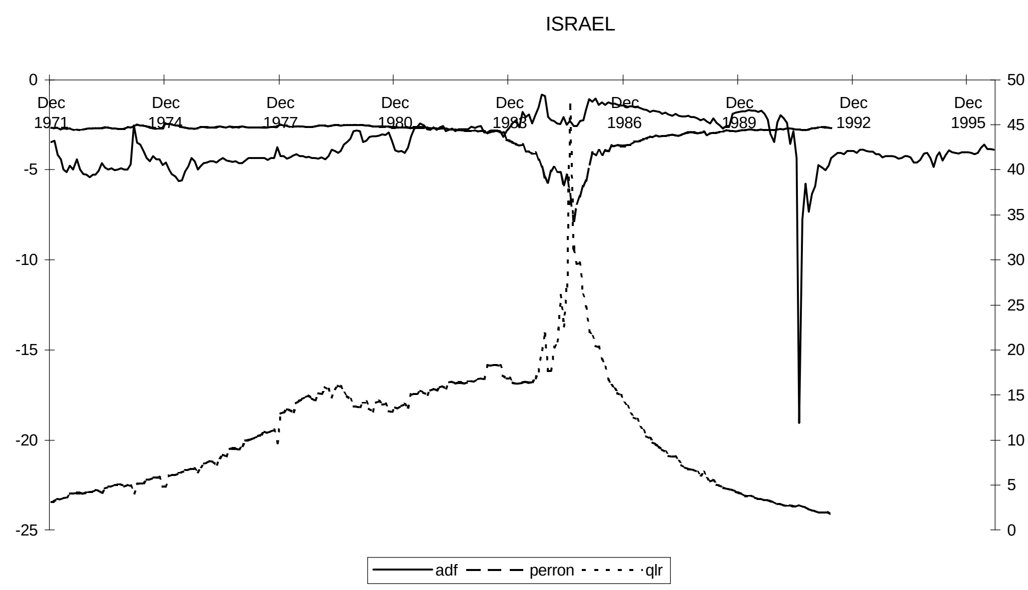
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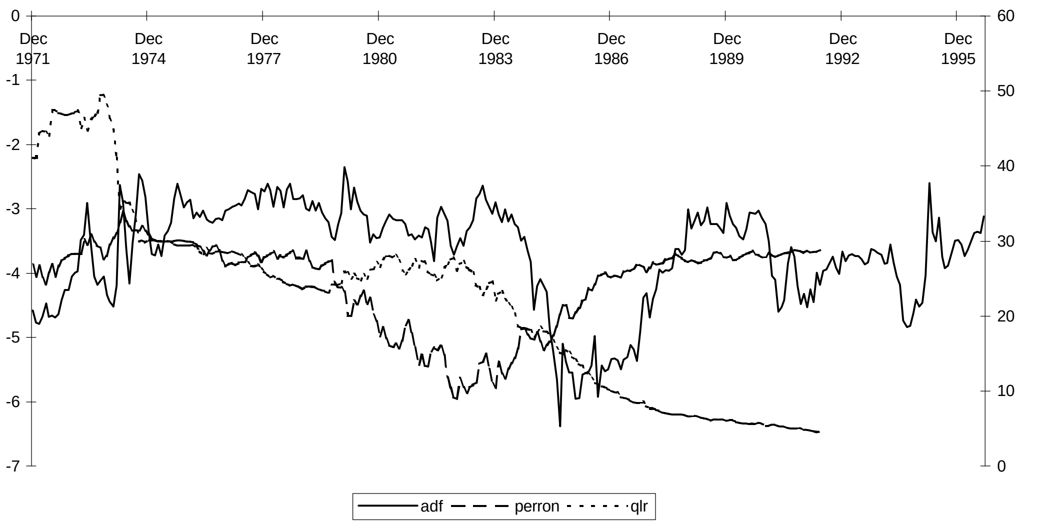

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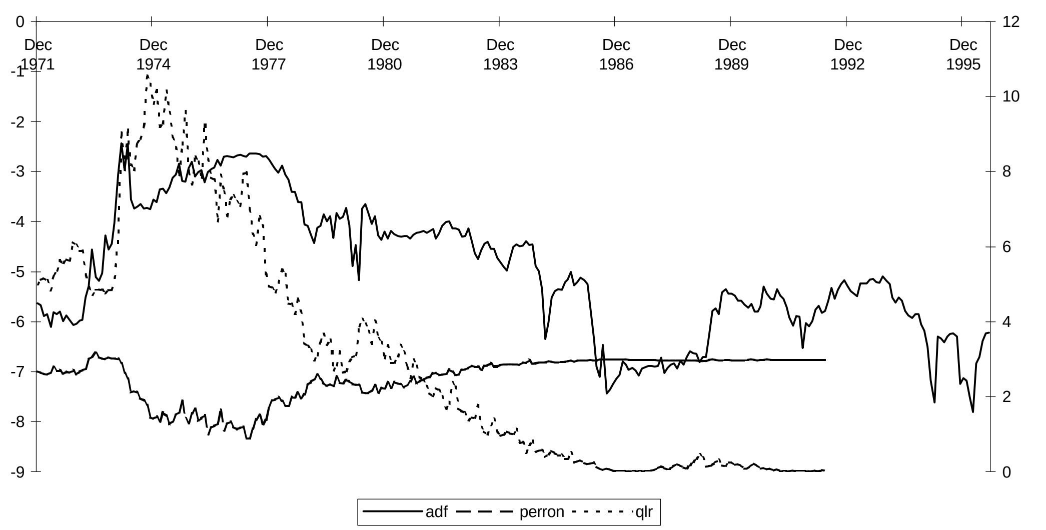

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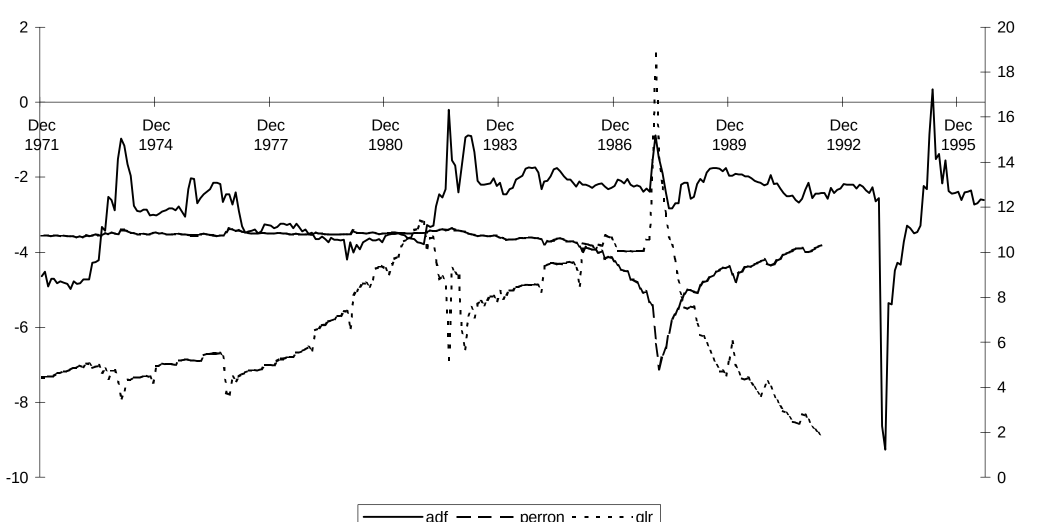

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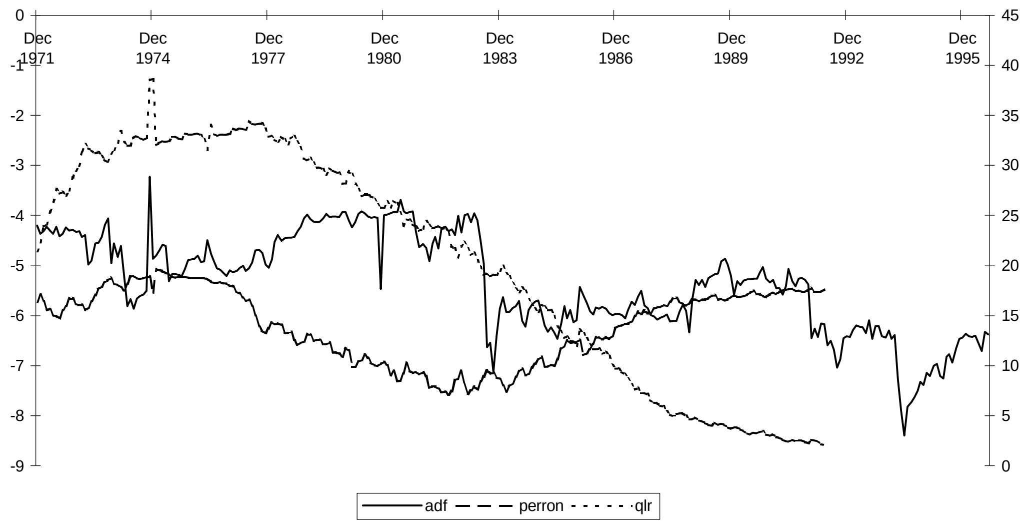

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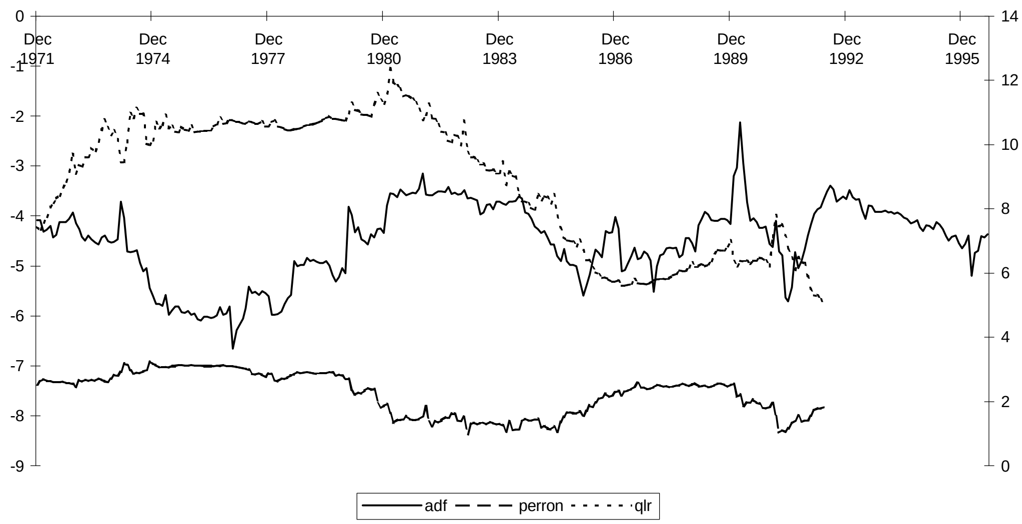

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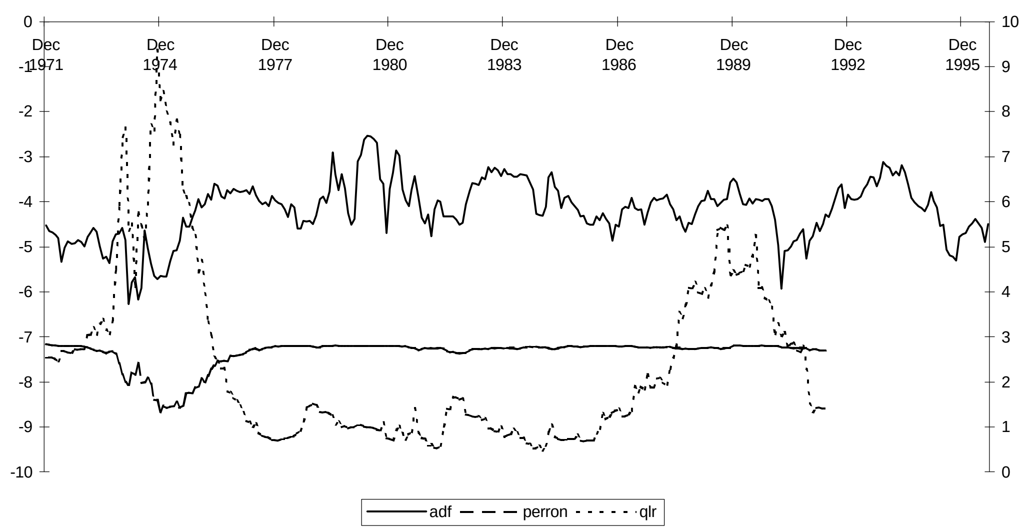

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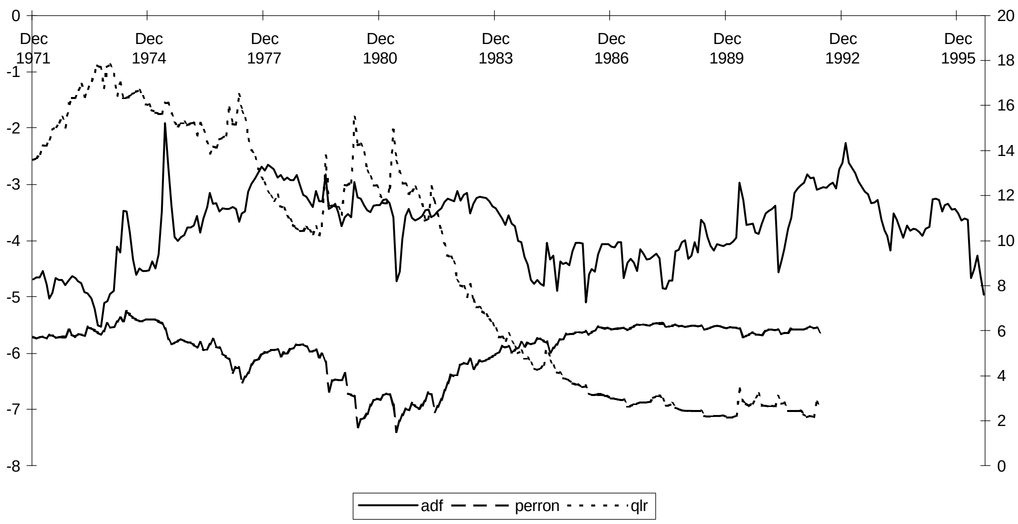

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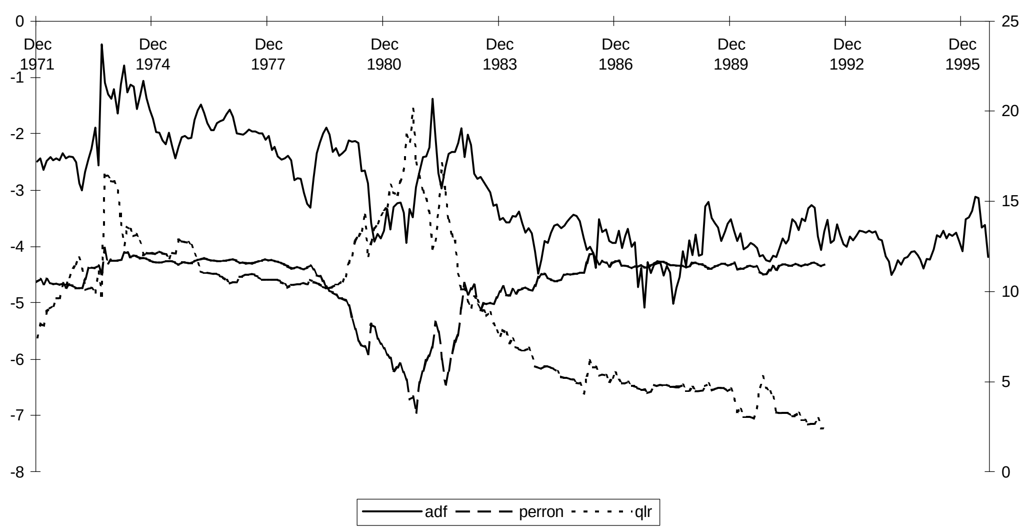

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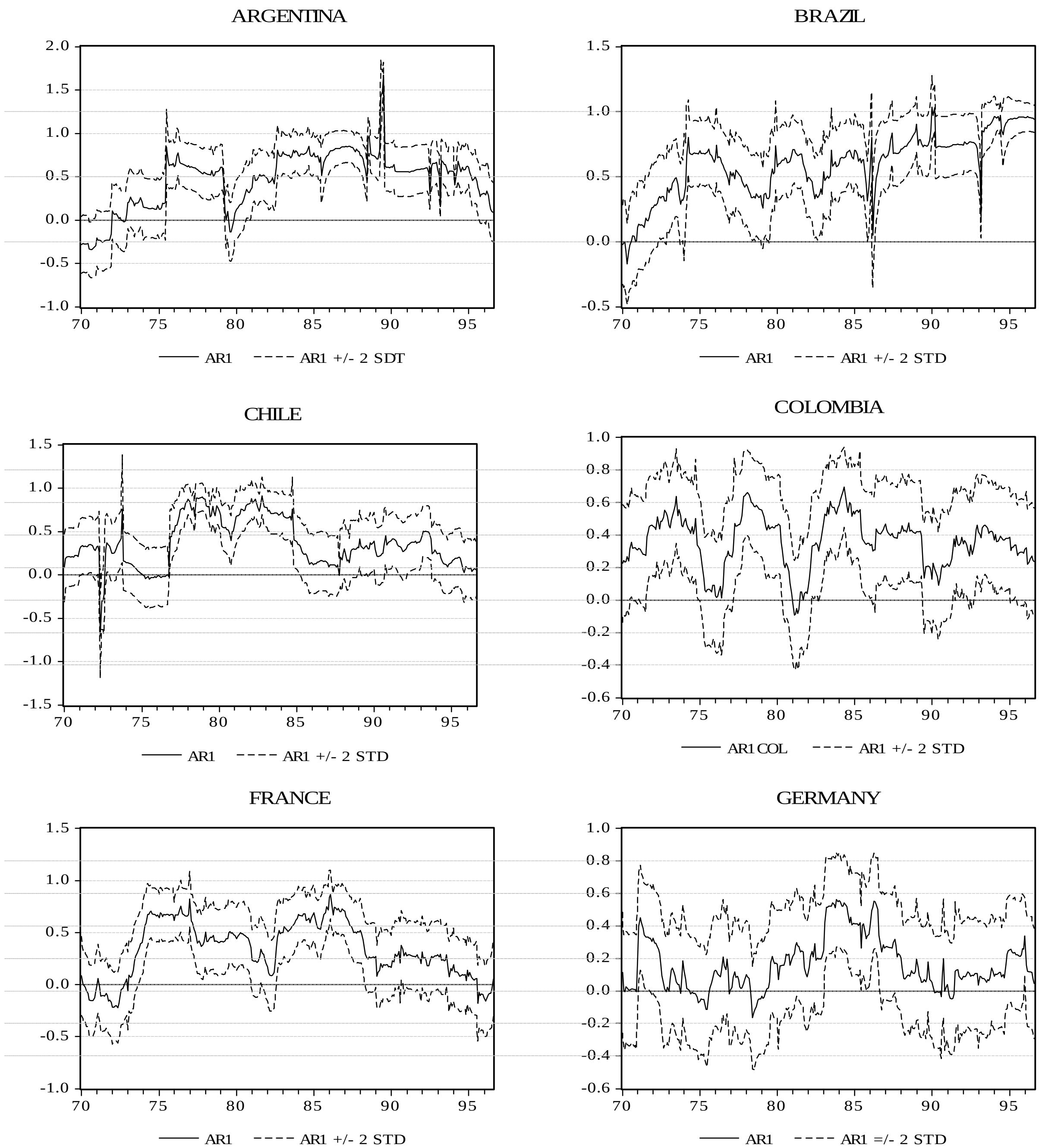
Figure 3. Rolling AR1



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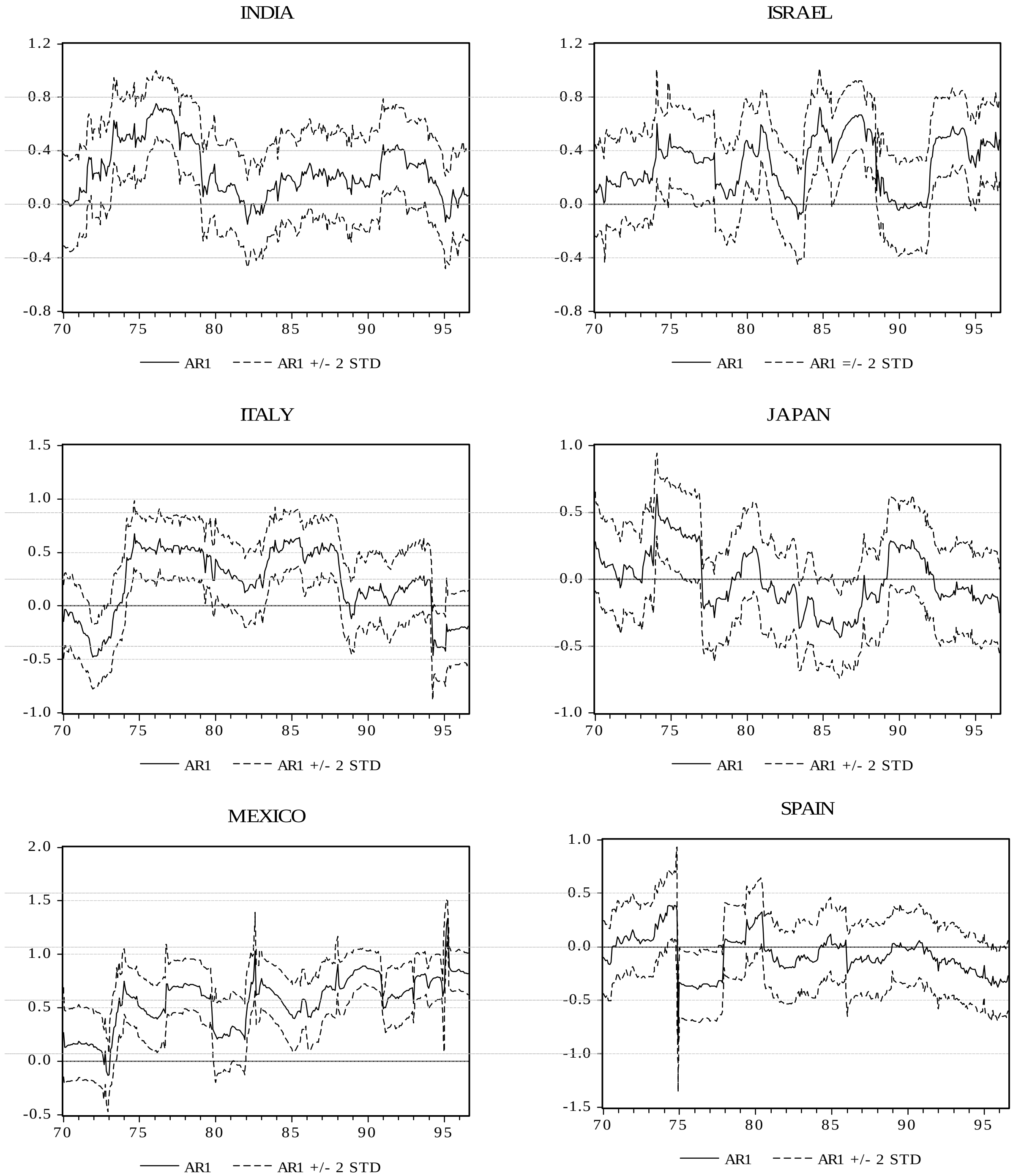


Figure 3. (continued)

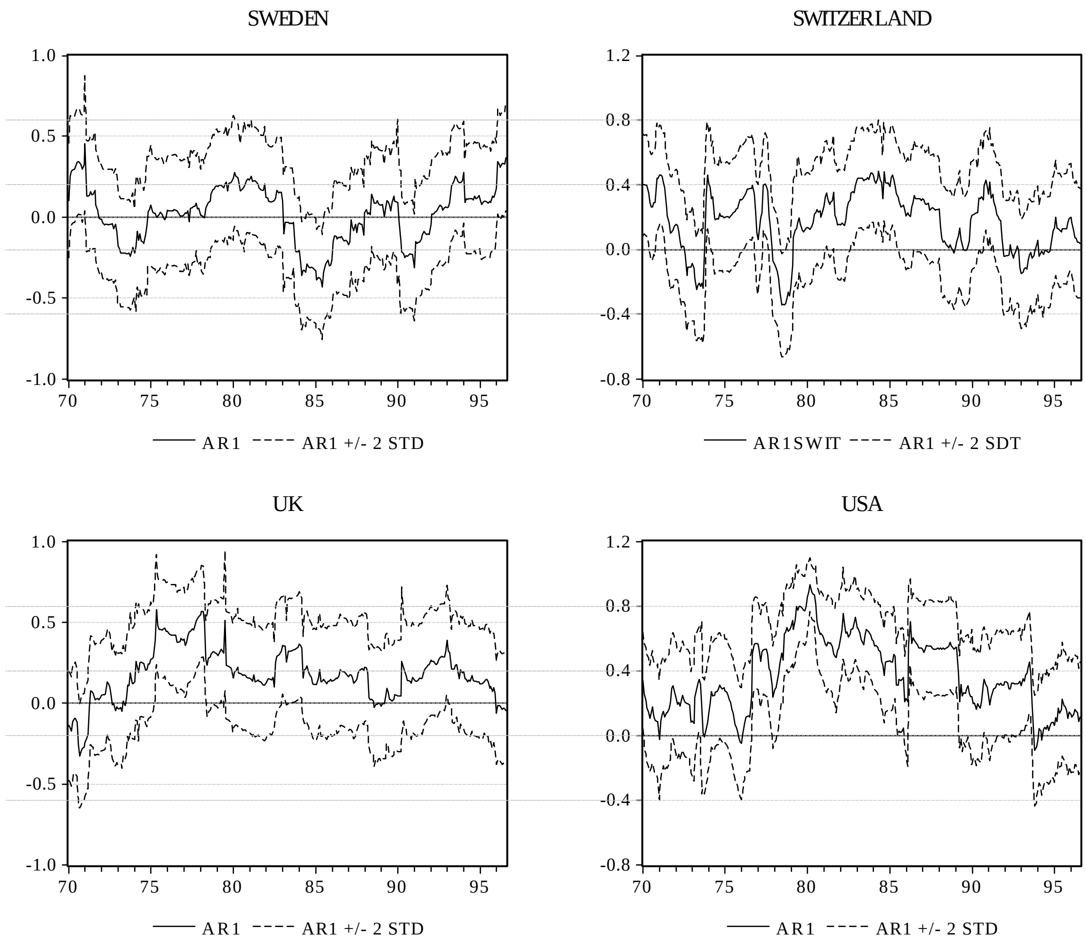
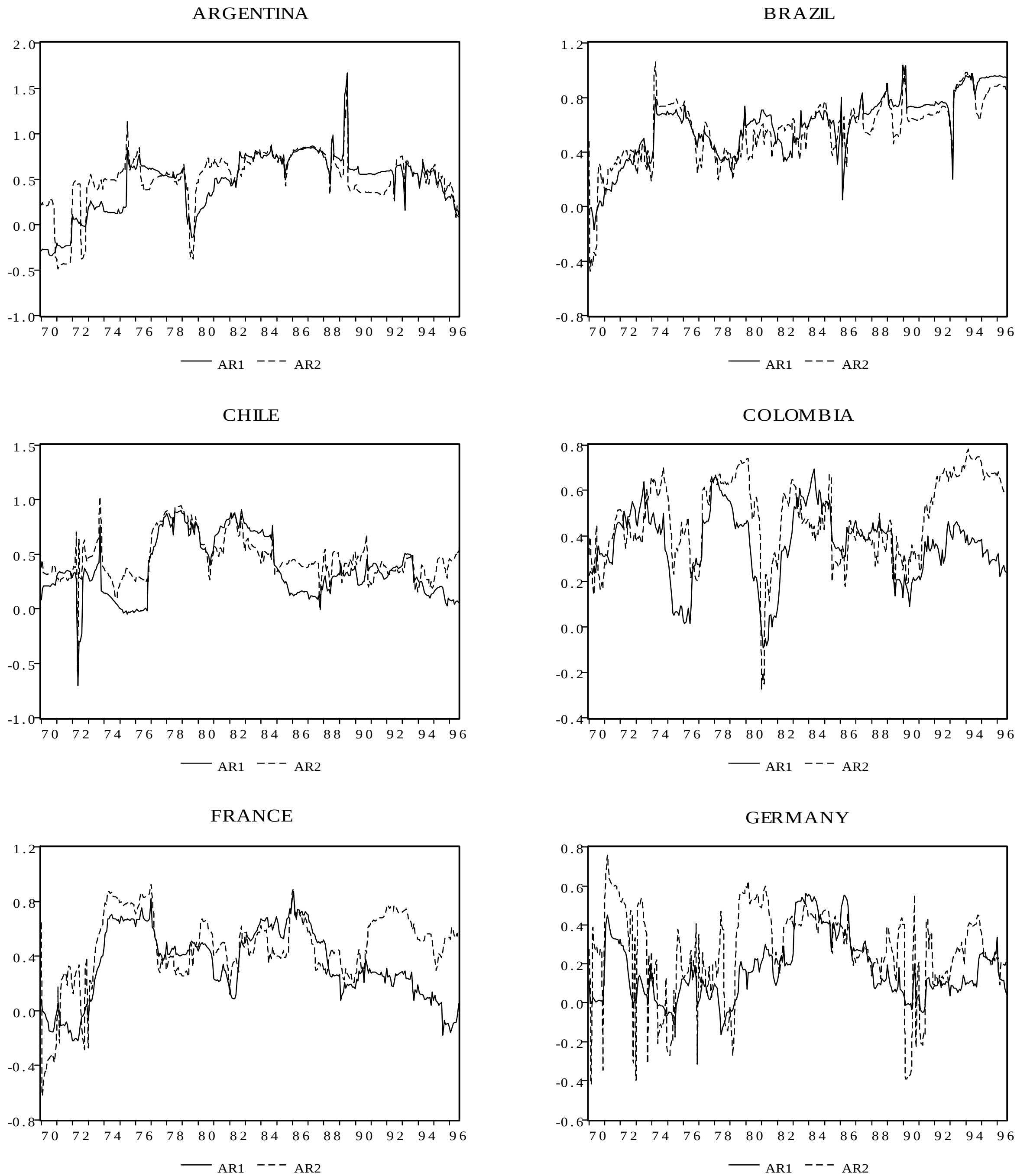


Figure 4. Rolling AR1 and AR2



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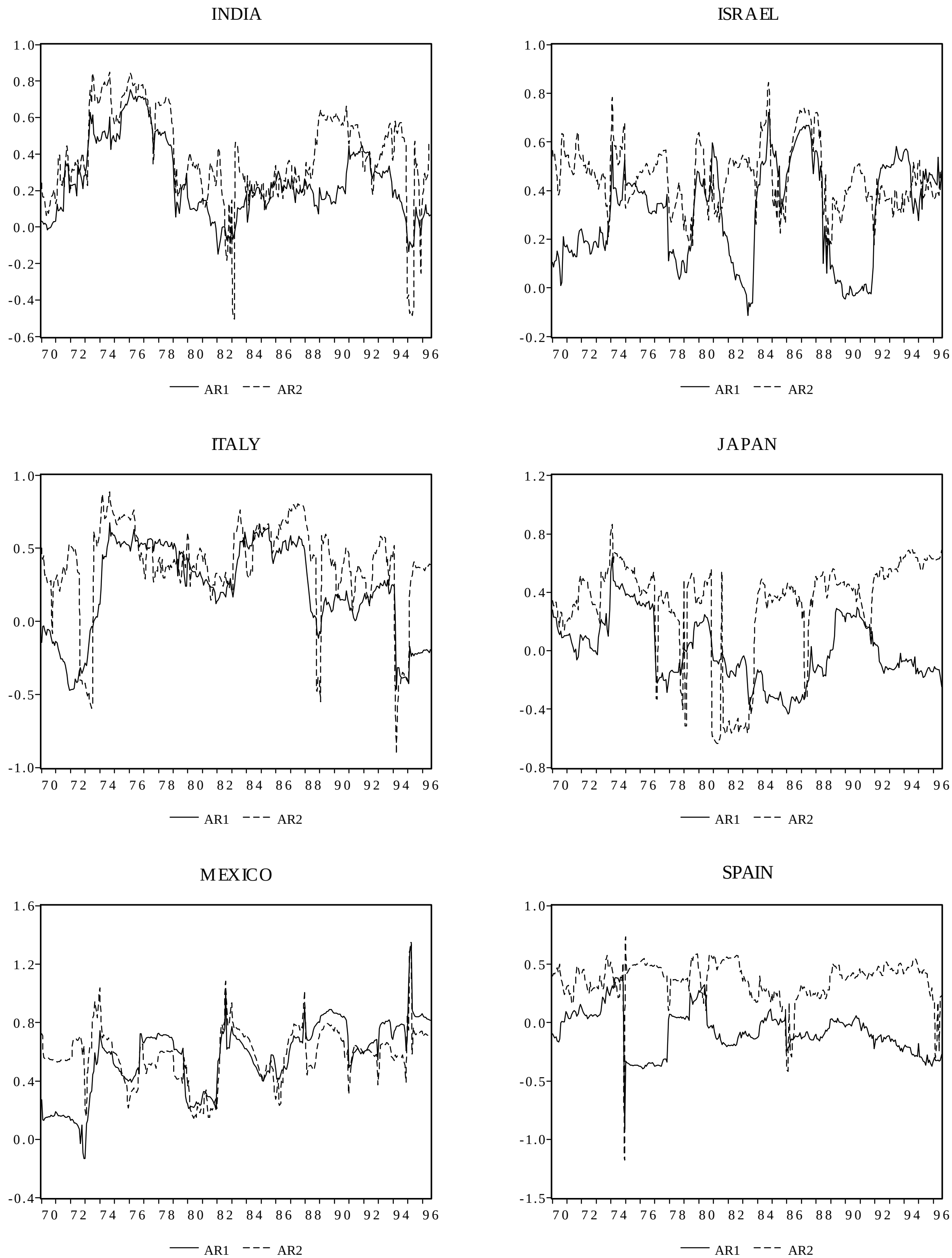


Figure 4. (continued)

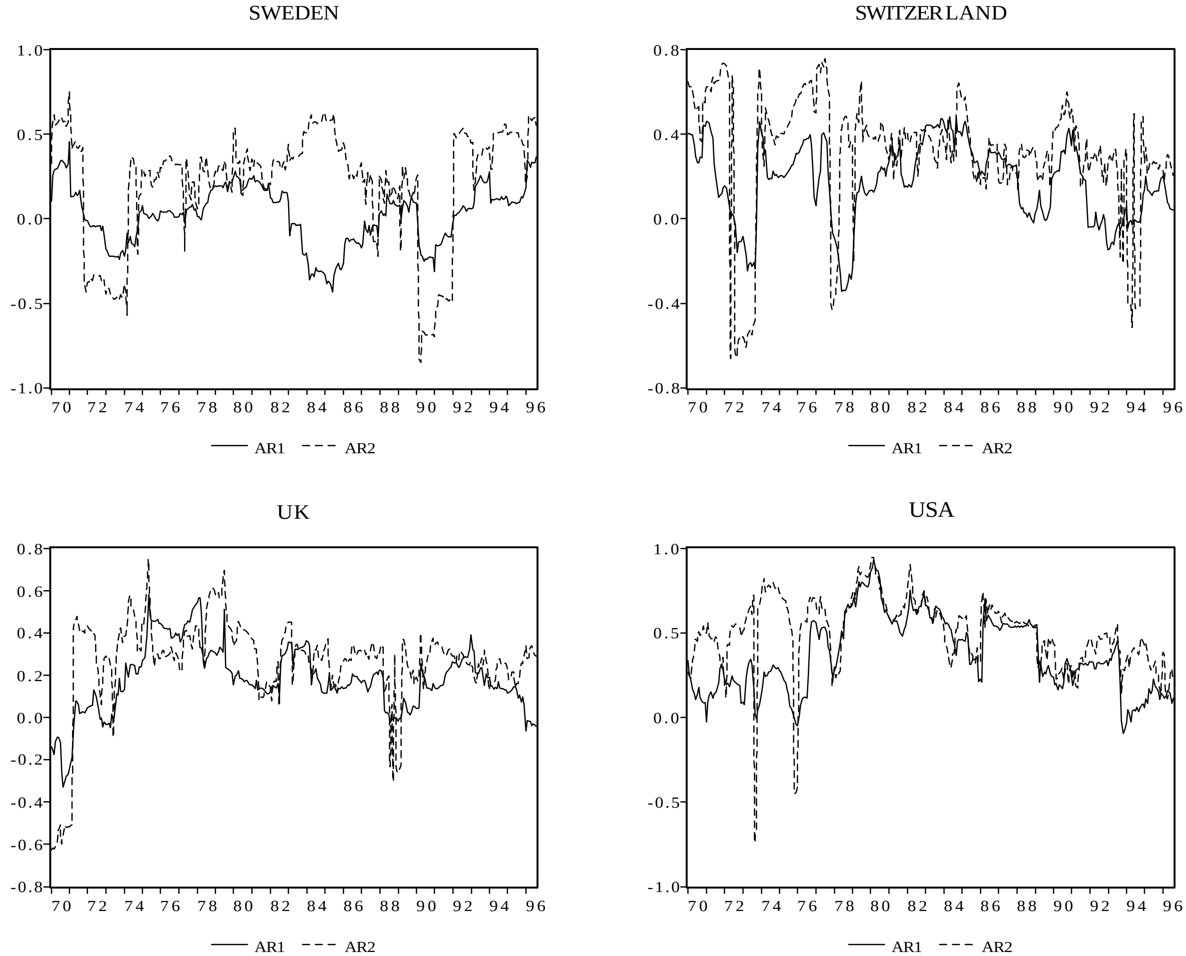


Figure 5. Rolling AR1 Including money growth

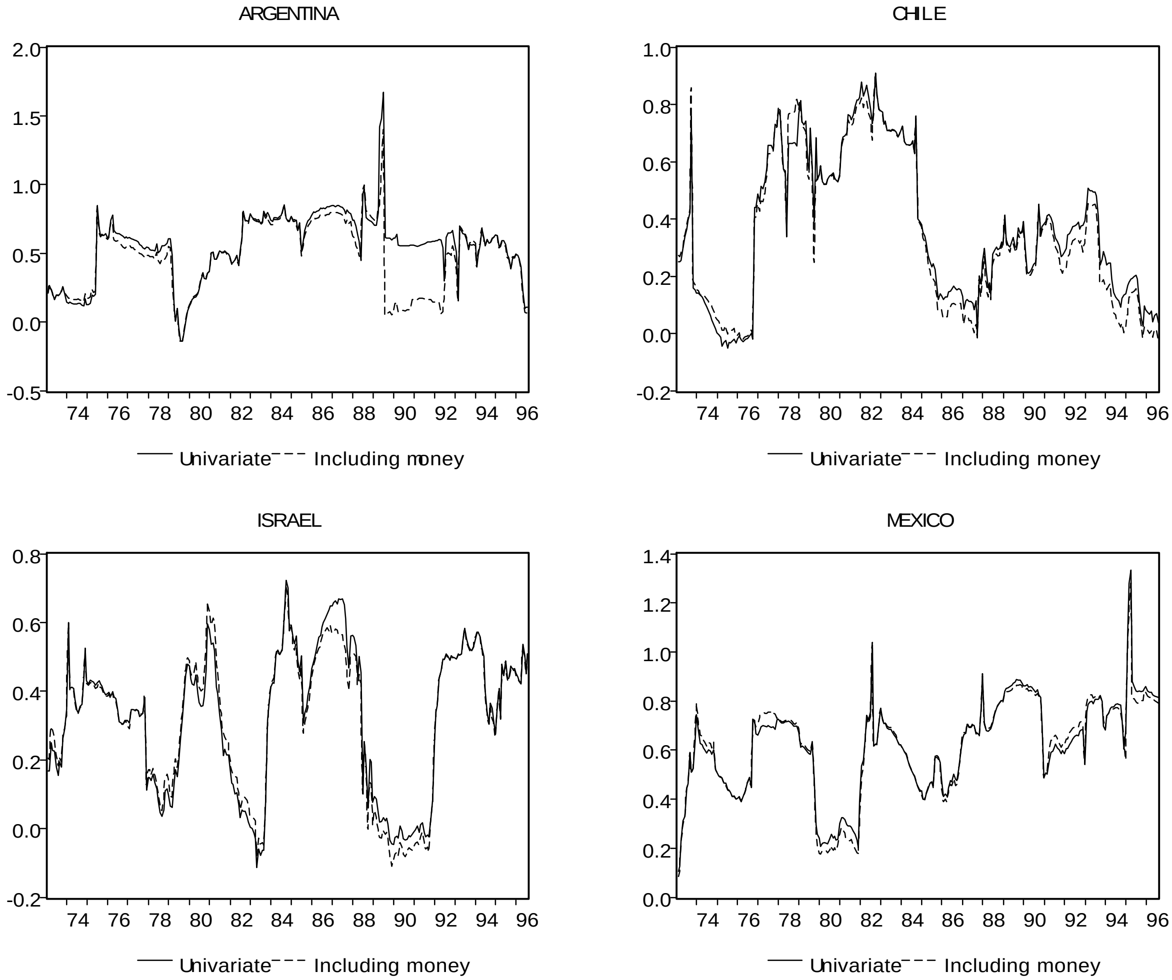


Figure 6. Inflation and Persistence

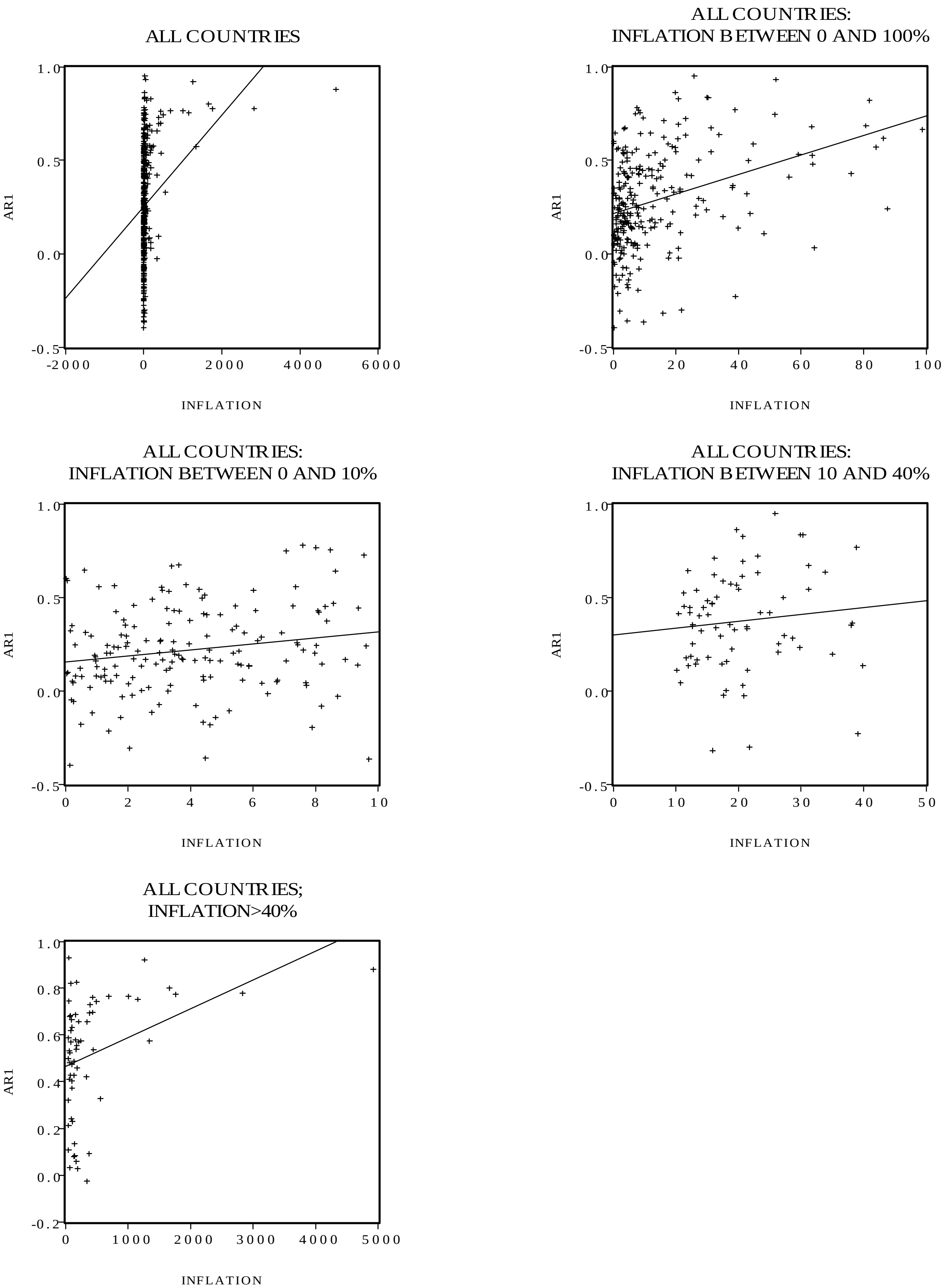


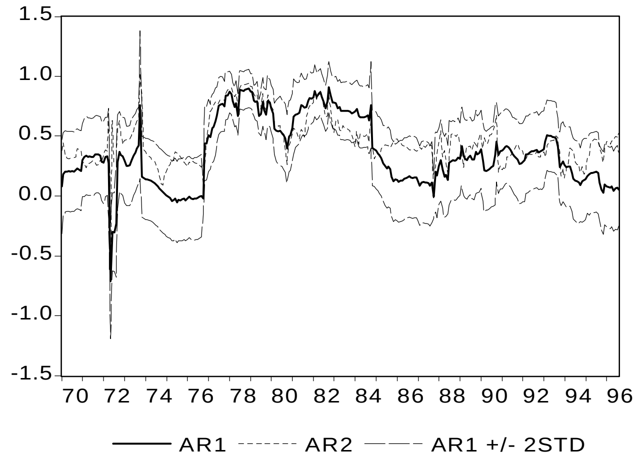
Table 5. Panel Regressions of Inflation on Persistence

(Standard errors in parenthesis)

	Full sample	p < 0.1	p < 0.4	p < 1	p >1
Pooled Least Squares	3.371 (0.552)	0.029 (0.009)	0.112 (0.017)	0.223 (0.030)	15.766 (4.897)
Fixed Effects	2.520 (0.695)		0.034 (0.015)		
Random Effects	2.825 (0.639)		0.037 (0.014)		
Number of observations	416	281	355	375	41

Figure 7. Inflationary Inertia: Chile and Mexico

CHILE



MEXICO

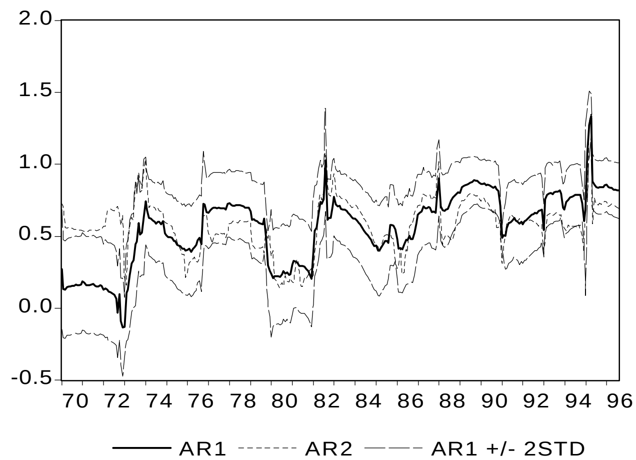
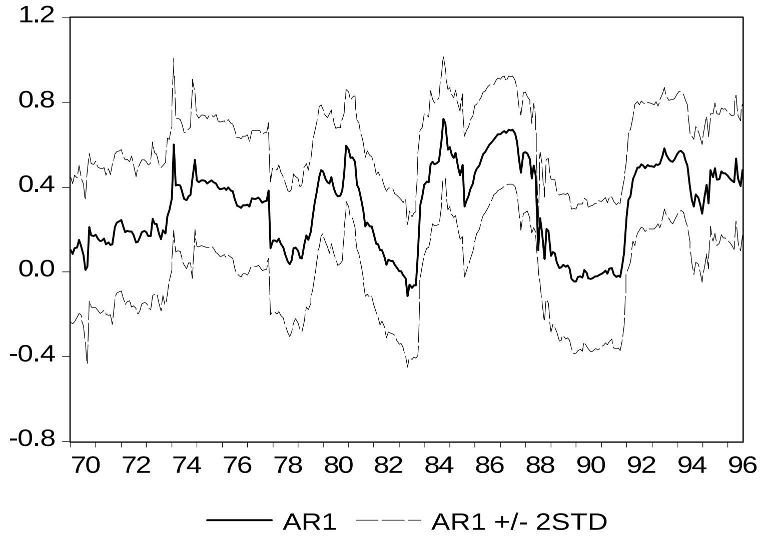


Figure 8. Israel's Exchange Rate Band

Figure 9. Inflationary Inertia: Israel



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